Pricing Government Contract Risk Premia: Evidence from the 2025 Federal Lease Terminations*

Soon Hyeok Choi†
Rochester Institute of Technology

Cameron LaPoint‡
Yale SOM

November 2025

[latest version here]

Abstract

Recent shifts in federal real estate policy, marked by lease cancellations and non-renewals, challenge the long-standing perception of government contracts as a secure and stable investment. We investigate how federal lease cancellations impact the pricing of government contract risk premia. Using unanticipated Department of Government Efficiency (DOGE) cancellations as a shock to commercial mortgage default risk, we find that first-loss CMBS bond tranches backed by loans tied to DOGE-notified leases experience a persistent 4% decline in price, with large, negative spillover effects to bond prices, delinquency rates, and rental cash flows tied to nearby private-tenant leases. These results reflect that early termination options were previously perceived by investors as a dormant clause. Applying arbitrage pricing models of commercial lease contingencies confirms the underpricing of risk associated with government tenants. Simulations of tail risks from early termination exposure result in aggregate property value losses in excess of $575 million for the Washington, D.C. securitized office market alone – well above total potential taxpayer savings from canceled lease payments.

Keywords: commercial real estate, government contracts, CMBS, lease contingencies, credit risk, arbitrage pricing

JEL Classifications: G18, G21, H83, R38

---

*We thank Natee Amornsiripanitch, Jiakai Chen, Andreas Christopoulos, Sean Flynn, Marco Giacoletti, Stefano Giglio, Paul Goldsmith-Pinkham, Robert Jarrow, Felipe Netto, Alex Priest, Franklin Qian, Tim Riddiough, Jacob Sagi, Nathan Seegert, Todd Sinai, and Alexei Tchistyi; as well as seminar and workshop participants at Yale, Northeastern University, and the UNC CREDA Real Estate Symposium for helpful comments and discussions. The Trepp CMBS/FEED, CoreLogic, Advan Weekly Patterns (c/o Dewey Data), and WRDS TRACE data used in this paper are licensed through Yale University Library. We thank representatives from Trepp, especially Ben Kater, for answering our questions about their data products. This research was financially supported by the Yale International Center for Finance and the Tobin Center for Economic Policy. We thank a diligent team of RAs for their excellent work, including Fernando Reyes De La Luz, Aurel Rochell, and Peter Williams. First draft: June 2025.

†Saunders College of Business, Rochester Institute of Technology. 107 Lomb Memorial Drive, Rochester, NY 14623. Email: **schoi@saunders.rit.edu**. Web: **steve-sh-choi.com**.

‡Yale School of Management. 165 Whitney Avenue, New Haven, CT 06511. Email: **cameron.lapoint@yale.edu**. Web: **cameronlapoint.com**.

1 Introduction

In early 2025, the newly created Department of Government Efficiency (DOGE) initiated the largest wave of lease terminations in federal history. Unlike standard federal contracts that include termination for convenience clauses, General Services Administration (GSA) leases operate under different regulations and rely on early termination options (ETOs) that can only be exercised during the "soft term" period of leases, typically after the initial five-year "firm term" with 90 to 120 days advance notice. DOGE's actions resulted in the cancellation of hundreds of leases within months, affecting nearly 9 million square feet of office space (Trepp Research, 2025) – a sharp contrast to the COVID-19 pandemic period when federal agencies largely maintained existing leases despite widespread remote work. This rapid and extensive use of previously dormant ETO provisions represented a fundamental shift in federal leasing practices and created significant disruptions to both government agency operations and the commercial real estate (CRE) market (Yahoo Finance, 2025).

We use the sudden onset of the DOGE lease cancellations and the large-scale nature of the shock to CRE markets as a natural experiment to estimate risk premia associated with government contracts.[1] In the immediate aftermath of the DOGE announcements, market observers speculated that the cancellations would ripple through U.S. commercial mortgage-backed securities (CMBS) markets, leading to increased default and loss risk (Bloomberg, 2025). We show that these fears were largely well-founded, with sharp cuts in prices commanded by the most junior CMBS tranches and declines in net operating income and debt coverage among directly impacted GSA-linked properties. We also find evidence of large negative spillovers to the performance of nearby non-GSA properties and their loans. Simulating the tail risks from ETO exposure across terminated, nearby non-terminated, and private-tenant leases, we find total property valuation losses for the Washington, D.C. market alone exceed taxpayer savings from canceled lease payments.

To motivate our empirical analysis, we extend the arbitrage pricing framework of Jarrow (2018, 2021) and Choi et al. (2025), applying contingency option pricing theory to CMBS backed by federal leases with ETOs. We compare two otherwise identical properties: one ignores the ETO, assuming zero probability of exercise, while the other accounts for its positive probability (ETO-salient). We show that treating the ETO as dormant leads to lower rents charged by landlords, indicating systematic underpricing of government contract risk. We quantify the resulting property-level value loss as the difference in risk-adjusted values under the two assumptions and derive a closed-form expression assuming independent Poisson processes

[1] As of February 2025, Fitch-rated CMBS bonds with GSA leases included as a top five tenant total $19.1 billion, across 280 loans and 344 properties (Fitch Ratings, 2025). Based on our calculations from GSA lease inventory data, as of December 2024, half of the federal government's 7,535 leases are up for renewal in the next five years, with the initial wave of cancellations leading to uncertainty about further fallout of the GSA-linked market.

for ETO notification and re-leasing. We also derive the insurance premium that would hedge against ETO risk.

Our model generates three key predictions. First, a more aggressive federal lease cancellation policy leads to greater expected property value losses, which translate to lower CMBS bond prices, reflecting a repricing of previously mispriced government contract risk premia and increased salience of ETOs. Second, properties that receive ETO notifications experience lower cash flows and values than comparable properties, due to premature terminations, higher vacancy risk, and reduced rental income. Finally, properties located near ETO-notified buildings exhibit declines in CMBS bond prices and net operating income, relative to otherwise comparable properties, reflecting significant negative spatial and capital market spillover effects from a market-wide wake-up call that also impacts private tenants.

We test these predictions using detailed data from Trepp combined with GSA lease inventory and hand-collected data from DOGE on lease cancellation event dates. We match properties across the datasets using multiple criteria, including address similarity, geospatial proximity, and exact matches on square footage and lease terms, with manual validation of potential matches. To confirm the plausibility of our match rates, we isolate the universe of commercial properties using property tax indicators in CoreLogic data and document that GSA properties are significantly overrepresented in CMBS pools relative to non-GSA properties, particularly in Washington, D.C. This new stylized fact underscores the importance of government contract risk to the broader CRE debt market.

Our empirical setting comes with several features that make it particularly well-suited for causal identification of government contract risk premia. Government termination actions occurred in quick succession, with most announcements concentrated on February 18, 2025 when the DOGE website first went live, and notification dates spanning a narrow window from January 30 to March 4, 2025. This concentrated timing around February 2025, combined with the substantial size of affected leases (averaging between 271,754 to 722,871 square feet across notification date waves), motivates our difference-in-differences (DiD) analysis. Our examination focuses on the Washington, D.C. MSA, which ranks first among metro areas in average square footage subject to early termination notifications while maintaining relatively high single-tenancy concentration, thus providing a compelling natural experiment to examine government contract risk repricing in CMBS markets.

Adopting a DiD research design that compares ETO-exercisable leases receiving termination notifications in the first wave of cancellations in February 2025 against similar soon-to-be ETO-eligible leases that did not receive notifications, we uncover evidence of immediate market adjustment to this policy shock. The results demonstrate economically sizable impacts across both financial and operational metrics. CMBS bond prices in first-loss group (FLG) tranches declined by 3.8% following ETO notifications, while property-level net operating income (NOI) fell by 5.6% relative to comparable untreated properties. These effects remain robust across multiple fixed effects specifications that control for location, temporal, structural, and bond-level characteristics. Due to the drop in NOI, debt service coverage ratios (DSCR) also fall, with the probability of covenant violations spiking in the aftermath of DOGE termination notices. The magnitude of these effects reflects the market's reassessment of risk that had previously been underpriced due to the perception of ETO clauses as legally available but operationally inactive.

We validate identification assumptions underlying the DiD approach by showing that there is no pre-treatment divergence in either bond prices, NOI, or DSCR, confirming that the policy shock to individual leases was likely unanticipated by market participants. On top of estimating no statistically significant pre-trends relative to DOGE notification, our results are robust to the pretest of pre-trends recommended by Roth (2022). Further, we allow for potential violations of parallel trends as in Rambachan and Roth (2023), but even under conservative assumptions about such violations we still obtain statistically significant effects of the DOGE actions on CRE markets. We also account for staggered cancellations and revoked cancellations in testing for whether there are asymmetric responses of the market to rescission decisions. For instance, the number of leases listed for cancellation by DOGE declined from a high of 793 in March 2025 to 384 as of September 30, 2025.

The cancellation of federal leases could also create negative spillovers to surrounding leases and properties through a combination of factors, including a drop in foot traffic to surrounding businesses due to a reallocation of federal workers and reduction in demand for the goods and services provided by local contractors to GSA tenants. Using standard spatial difference-in-differences approaches (e.g., Gupta et al., 2022; Chen et al., 2024) comparing properties linked to private-tenant leases in properties close vs. far away from a DOGE-terminated lease, we document a 10.1% drop in NOI backed by nearby properties with non-federal tenants. Decomposing this spillover effect, we do not find evidence that it is driven by contagion effects transmitting to non-ETO exercised leases packaged within the same bond pool alongside a DOGE-notified lease. This result holds even after controlling for local neighborhood time trends at various levels of geographic granularity, helping us combat a common challenge in such "ring" research designs in which selection into treatment status depends on distance cutoffs which may simply reflect other secular changes to neighborhoods rather than the local event of interest (Diamond and McQuade, 2019; LaPoint, 2022).

Feeding in our reduced form estimates of the decline in NOI from GSA and non-GSA leases stemming from early termination risk exposure, we simulate implied property valuation losses from our arbitrage pricing framework for the securitized portion of the Washington, D.C. CRE market. We compute two measures to capture tail risk: the minimum loss incurred in the worst 5% of simulated outcomes (i.e., the 95% Value at Risk) and the average loss conditional on being in the worst 5% of outcomes (i.e., the expected shortfall). We calculate property value destruction, in a 95% Value at Risk sense over a five-year period, of $914 million for ETO-eligible lease properties, $256 million for DOGE-notified lease properties, and $298 million for non-GSA properties. In the median scenario, the combined total expected shortfall is $575 million, implying an additional $51 million loss in terms of five-year local property tax revenues. These value losses are close to the entire savings from canceled lease payments nationwide and dwarf savings from the D.C. metro area of between $75 million and $100 million reported during the peak termination period in March 2025. Hence, the sudden shift in federal real estate policy generates potentially larger private-sector value erosion than it does taxpayer savings.

We contribute to several literatures at the intersection of real estate, finance, and political economy. The commercial real estate literature does not isolate lease contingency clauses – such as the ETOs we study – as a distinct driver of CMBS or commercial debt pricing, and any lease-related effects appear only as part of broader risk and contract structure analyses. For instance, Mooradian and Yang (2000) document in a small sample of commercial leases that tenants who select into leases with a downsizing option tend to pay higher rents. Similar to tenant exercises of downsizing options, Glancy and Wang (2023) show that even prior to the pandemic, lease expirations increase downside risk to a property's occupancy and income. Cheng et al. (2022) study how lease structures, including the negotiation of contingency clauses, change following the introduction of a new lease accounting rule (ASC 842) in private debt markets.

With the exception of Allen et al. (1997), who compare federal government to private-market office lease pricing in two U.S. states, there is virtually no empirical work isolating the GSA lease segment of the market. Studying government agency leases is important in its own right given the large average size of the leases, large dollar values of attached loans, and the potential for spillovers to local economies through employment and foot traffic.[2]

Prior work in urban economics underscores the special role of anchor tenants in the retail sector on lease (Pashigian and Gould, 1998; Choi et al., 2025) and product pricing (Konishi and Sandfort, 2003). Spatial spillovers of underperformance in real estate markets arise due to hyperlocal agglomeration economies in CRE (Rosenthal and Strange, 2020), with consumption externalities spread across retail tenants due to shoppers following travel itineraries to run errands between home and work (Miyauchi et al., 2025). The failure of anchor tenants can also lead to agglomeration of bankruptcy cases (Benmelech et al., 2018). However, even conditional on the within-submarket geographic proximity of properties to a terminated lease, local foot traffic, and other neighborhood business health metrics, we conclude CMBS repricing due to government policy risk can propagate through capital markets to nearby non-GSA-linked leases. Indeed, we find no discernible changes in foot traffic to retail or non-retail establishments located nearby a DOGE-canceled office lease, casting doubt on the importance of consumption externalities for the valuation multiplier we uncover in our setting.

By studying a new natural experiment consisting of realized political risk to the real estate sector, we contribute to work quantifying adverse economic consequences of political

[2]According to the Office of Personnel and Management's (OPM) FedScope data, 2.3 million workers, accounting for roughly 2% of the U.S. civilian workforce, were employed across all federal agencies as of September 2024. See `https://www.fedscope.opm.gov/` for the most current snapshot of federal employment counts

uncertainty. This literature has largely focused on government shutdowns, or the threat of shutdowns. Baker and Yannelis (2017) report a marked decline in household consumption during the 2013 shutdown. Gelman et al. (2020) estimate a spending reduction of 58 cents per dollar lost in liquidity during the same period. Herpfer et al. (2023) describe declines in government productivity – such as reduced accounting processing and patent creation – that persist up to four years post-shutdown. Baker et al. (2016) show more generally beyond shutdowns that policy uncertainty indexed throughout a century of U.S. newspaper data predicts declines in real aggregate activity and other major economies, echoing previous findings of delayed corporate investment spending during election years (Julio and Yook, 2012). For the purpose of quantifying government risk premia, our empirical setting has an advantage of clear "treatment" dates relative to the lengthy disputes related to Congressional budget appropriations or election-related uncertainty,[3] in the sense that individual government contracts are directly notified of termination on known dates, while other leases are not.

A large finance literature estimates how investors price political risk into securities markets. Pástor and Veronesi (2012) and Pástor and Veronesi (2013), respectively, show through the lens of general equilibrium models how stock prices negatively react to policy announcements and policy uncertainty generates a sizable risk premium. Consistent with these predictions, Brogaard and Detzel (2015) conclude that the economic policy uncertainty index of Baker et al. (2016) positively forecasts short-term abnormal returns on equities. Several empirical studies report that election uncertainty lowers liquidity and trading volume (Pasquariello and Zafeiridou, 2014) as well as firm valuations and productivity (Col et al., 2017), while raising option-implied volatility and hedging costs (Goodell and Vähämaa, 2013; Saiegh, 2023). Hassan et al. (2019) use firm-level textual analysis to show that higher political risk coincides with increased stock return volatility, a rise in risk dispersion, and a large uptick in lobbying activity. Kelly et al. (2016) emphasize that equity options spanning political events tend to be more expensive, as they offer a hedge against political tail risks. Similarly, our results indicate that the implicit put options represented by federal ETOs which are currently vested are cheaper because of the heightened loss exposure they carry.

A key distinction of our paper vis-à-vis extant research on the financial pricing of political uncertainty is our analysis of debt markets rather than equity markets. Moreover, relative to settings such as Congressional budget debates or contentious elections, the temporal clustering of government termination actions and the absence of pre-trends support our causal interpretation that DOGE's intervention served as a "wake-up call," revealing latent exposure in CMBS structures to federal lease terminations. Our results highlight how previously dormant contractual provisions can become salient sources of credit risk once activated, demonstrating the need for more explicit pricing of government contract risk in securitized products.

[3]Prior to the Fall 2025 shutdown episode which lasted 43 days, the three longest shutdown events in U.S. history lasted 35 days (2018–2019), 21 days (1995–1996), and 16 days (2013). Several weeks of political gridlock over budget reconciliation preceded each of these events.

Finally, the negative consequences for debt markets we document from the regime shift in government lease contracts may magnify risk exposure of regional banks following the collapse of office real estate values due to the transition towards remote work after the pandemic (Gupta et al., 2025). Indeed, our simulation exercises point to a lower bound for property value destruction from early termination exposure of nearly 4% of aggregate office market values in Washington, D.C. Jiang et al. (2025) simulate risks of insolvency for up to 300 banks due to losses on CRE loans amplified by the sharp monetary policy tightening cycle between 2022 and 2023 (Jiang et al., 2024). On top of regional banks, Brown et al. (2024) argue that major life insurance companies hold as much as 16% of their portfolio in CRE debt, of which roughly one-third includes CMBS investments, with GSA-concentrated Washington, D.C. ranked third in terms of the underlying property location of life insurers' book value mortgage exposure. Such trends may lead banks most exposed to CRE debt to reduce their credit supply to non-CRE investments, thus hampering economic growth in other sectors (Anenberg et al., 2025).

The rest of this paper is organized as follows. Section 2 provides background on the 2025 federal lease terminations we explore in our empirical setting. Section 3 develops an arbitrage pricing framework for valuing lease contingencies in the CMBS market. Section 4 describes our commercial property market data and sample construction. Section 5 outlines our empirical strategies. Section 6 presents our main empirical results. Section 7 uses our empirical estimates to simulate implied property valuation losses. Section 8 concludes.

2 Background on Federal Lease Terminations

In early 2025, federal leasing policy underwent a significant transformation, marked by a sharp increase in the termination of leases administered by the General Services Administration (GSA). In contrast to standard federal procurement contracts governed by the Federal Acquisition Regulation (FAR), which often incorporate termination for convenience clauses, such as FAR 52.249-2, GSA leases are governed by the General Services Administration Acquisition Regulation (GSAR) and generally do not include such provisions (Federal Acquisition Regulation, 2025). Instead, the GSA can exercise *an early lease termination option* (ETO) within a pre-defined period of the lease term. In the post-2015 GSA lease inventory, every lease contract has a termination right date.

As shown in Figure 1, GSA leases are typically structured into two sequential phases. The *firm term*, usually encompassing the first ten years, prohibits early termination and provides cash flow certainty for lessors. The subsequent *soft term*, typically five years in length, allows the GSA to unilaterally terminate the lease provided that it issues an advance notice of 90 to 120 days (U.S. General Services Administration, 2023, 2024). Consequently, among contracts that were subject to the 2025 wave of federal lease cancellations, 85% had already transitioned into the soft term. In nearly all of the remaining 15% of cases the agency's operations were either shut down or the agency approved of the termination in an effort to downsize.[4] Figure 3 provides examples of the termination right clauses in the standardized lease contract (form L100) that GSA uses when negotiating with private landlords.

Leases announced as terminated by DOGE are geographically dispersed, with each state having at least two federal leases canceled (Figure 4). However, the square footage of terminated leases is highly concentrated in a small handful of areas, particularly Washington, D.C. and Atlanta, Georgia (Figures 5 and 6). For single-tenant leases, termination counts are greatest in larger states (Figure 7). Savings, as measured by total rent payments which will no longer be paid, are more geographically dispersed due to regional differences in per square foot rents (Figure 8).[5] Tenants of terminated leases span nearly one-hundred different federal agencies.

The scale of lease terminations initiated in 2025 is unprecedented in federal leasing history. Based on historical GSA inventory lists available from 2015, one-year GSA lease termination rates fluctuated between 2% and 3% before spiking above 5% in the first half of 2025 (Figure 9). The spike in termination rates is even more pronounced when computed over the subset of GSA leases which are in the soft term, whereby the lease is eligible for early termination in the absence of force majeure or agency closure; for soft term leases, one-year cancellation rates spiked from their historical 3-4% average to 12% in March 2025 (Figure 10).[6]

The jump in termination rates for ETO-eligible leases supports the notion that the cancellation option embedded in government leases would have been viewed by market participants as a rarely exercised, dormant clause. Notably, during the COVID-19 pandemic, despite widespread remote work, federal agencies largely maintained existing leases, citing the absence of cancellation clauses and potential costs of early termination. In contrast, the 2025 termination involved hundreds of leases spanning millions of square feet of office space and widespread disruption of federal agency operations. The rapid pace and breadth of terminations marked a significant departure from prior federal leasing practices.

Figure 11 shows how the number of leases slated for termination by DOGE evolved since the DOGE website's publication on February 18, 2025. The number of canceled leases peaked at 793 between March 13, 2025 and March 18, 2025, dropping down to 384 leases as of September 30, 2025. These cancellations are relative to an inventory of 7,535 GSA leases as of December

[4]In two-thirds of all terminated leases, DOGE describes the termination on their website as being conducted via “mass modification," a method for modifying lease contracts under the GSA's Multiple Award Schedule (MAS) program, whereby government agencies receive services at pre-negotiated prices from private contractors.

[5]DOGE calculates a contract's value using a federal contracting concept known as “total potential value," which includes any remaining lease payments left on the current term of the lease, plus any payments in the event all remaining lease renewal options are exercised (CBS News, 2025).

[6]These calculations assume that DOGE announcements eventually lead to removal of the lease from the GSA inventory list after the 90 to 120 day grace period for ETO-exercised leases. As shown in Figure 11, some leases were eventually removed from the DOGE savings webpage, leading to ambiguity about whether such a lease will eventually be removed from the government's inventory. Therefore, our series captures the perceived spike in cancellation rates as of 2025Q1 but not necessarily finalized terminations.

2024.[7] Many of the leases which were at one point listed as terminated on the DOGE savings website but then subsequently removed either remain vacant or have since been leased out to a new, non-governmental tenant.

Still, the presence of both cancellation and rescission announcements motivates staggered versions of our standard difference-in-differences design, discussed in Section 5, in which the treated status of properties and leases turns on and off over time (de Chaisemartin and D'Haultfoeuille, 2020). The ebb and flow of DOGE decisions also allows us to test for effects on the CMBS market from moving between regimes with higher vs. lower lease cancellation risk. In the next section, we theoretically model such transitions and derive testable implications for commercial real estate.

3 Arbitrage Pricing Framework for CMBS Market

Our theoretical framework extends the contingency option pricing model introduced by Choi et al. (2025). First, we apply the arbitrage pricing model to two otherwise identical properties that differ solely in their treatment of an ETO: one property presumes the ETO will not be exercised, assigning it zero probability, whereas the other assigns a strictly positive probability to its exercise. We show that the property which treats the ETO exercise as impossible exhibits lower rent payments, reflecting an underpricing of the government contract risk premium. Furthermore, we define the loss from mispricing the ETO as the difference in values between the two properties, deriving a closed-form solution under the assumption of Poisson arrivals for both ETO notification and re-leasing processes. Finally, we formalize a federal lease contract without embedded options as analogous to a coupon-bearing bond; the landlord effectively lends commercial space, representing the bond principal, to the federal tenant, who utilizes the space for operations and returns it at lease maturity. The tenant's periodic rental payments correspond to the bond's coupon payments. If the lease with an embedded ETO correctly priced the government contract risk premium, the federal tenant would additionally compensate the landlord through an insurance premium.

3.1 Model Setup

We adopt the arbitrage pricing framework developed in Jarrow (2018, 2021) and Choi et al. (2025) to evaluate commercial lease contingencies. Given that federal leases have finite terms, we assume a continuous trading model with a finite horizon T^*. We formalize the characterization of uncertainty in the model with a filtered probability space $(\Omega, \mathscr{F}, \mathcal{F} = (\mathscr{F}_t)_{t\in[0,T^*]}, \mathbb{P})$ where

[7] The DOGE website does not provide a unique identifier to track leases over time. We create a lease panel identifier based on a combination of the city-level location, government agency tenant, and square footage. We do this after correcting for misspellings/abbreviations in the agency name after cross-referencing with the historical GSA lease inventory list (e.g., "EPA" vs. "Environmental Protection Agency").

Ω is the state space, $\mathscr{F}$ is a σ-algebra of events, $\mathcal{F} = (\mathscr{F}_t)_{t\in[0,T^*]}$ is an information filtration, and $\mathbb{P}$ is the statistical probability measure. The traded assets are a default-free money market account and default-free zero-coupon bonds. We denote the time t price of a zero-coupon bond maturing at T and always paying a dollar as $p(t,T)$, and r_t is the default-free spot rate of interest. Define the time t value of a money market account as

$$B(t) = e^{\int_0^t r_u du}$$

Since we assume the market to be arbitrage-free, the First and Third Fundamental Theorems of Asset Pricing are satisfied. The theorems imply that there is an equivalent probability (i.e., risk-netural probability) measure $\mathbb{Q}$ with respect to $\mathbb{P}$ such that

$$p(t,\mathbb{T})e^{-\int_0^t r_u du} \tag{3.1}$$

is a martingale for all $\mathbb{T} \in [0,T]$.[8] We model the exercise of an ETO by a federal tenant as a stopping time $\tau \in [0,T]$ with respect to the underlying information filtration, where $T < T^*$, and T denotes the original lease expiration date.[9]

The primary credit event from the landlord's perspective, and thus for the CMBS pool and bondholders as well, occurs when the federal tenant elects to exercise its ETO and issues formal notice to the landlord in accordance with the lease's prescribed advance notice period. Figure 2 illustrates the sequence of key events that unfold following an exercised ETO. The process begins at time τ, which represents the moment (stopping time) the tenant delivers formal notice of its intent to terminate the lease early. This notice must be issued in accordance with the lease's advance notice requirement denoted by a fixed period α, typically ranging from 90 to 120 days. During this notice window, the federal tenant remains contractually obligated to pay rent, allowing the landlord a limited time frame to begin preparing for the impending vacancy.[10]

Following the expiration of the advance notice period, the lease is effectively terminated, and the federal tenant vacates the premises. At this point, the landlord becomes responsible for the contract rent associated with the now-vacant space. This rental loss continues until a new tenant is secured at a future date (stopping time), denoted by η. The interim period between lease termination and tenant replacement exposes the property owner to a direct decline in net operating income (NOI).

---

[8]The formal notion of an arbitrage-free market is known as a market that satisfies the principle of *No Free Lunch with Vanishing Risk* (Jarrow, 2018).

[9]To accommodate specific times of key events in an exercised ETO episode, we equate the finite time period $[0,T]$ with $[t_0,t_m]$ for $m \in \mathbb{N}$. For the time t valuation, it implies the fixed horizon is $[t_j,T]$ for $j \in \mathbb{N}$.

[10]While the original lease contract typically specifies the grace period, additional rent abatement may apply if a federal tenant vacates the premises before the grace period ends, as illustrated by the "vacant premises" clause in the standard GSA leasing form in Panel B of Figure 3. Hence, our baseline model assumes $\alpha = 0$ for simplicity.

3.2 Rental Cash Flows

Consider two otherwise identical properties that differ only in the salience of the ETO exercise probability: one property assumes the early termination option (ETO) exercise event has zero probability while the other assumes a strictly positive probability of exercise. The no-arbitrage market value of the property in which the ETO is assumed to have zero probability is:[11]

$$\frac{V(t)}{B(t)} = \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{T} \frac{R}{B(s)} + \frac{V(T)}{B(T)}\right] \tag{3.2}$$

Intuitively, the current discounted value of the property with a dormant ETO equals the risk-adjusted discounted sum of all future rent payments plus the terminal liquidation value of the property. Since the ETO is considered a zero-probability event, there is no lapse of lease payments during the contract period.

Consider another property with a strictly positive probability of ETO exercise. This introduces a friction whereby space rendered vacant by a prematurely terminated lease cannot be leased out again until a certain period, $\eta > \tau$, has passed. This implies that the ETO-salient property's market value is:

$$\frac{\widetilde{V}(t)}{B(t)} = \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{\tau} \frac{\widetilde{R}}{B(s)} + \sum_{s=\eta+1}^{T} \frac{\widetilde{R}}{B(s)} + \frac{V(T)}{B(T)}\right] \tag{3.3}$$

The next proposition characterizes how correctly pricing a dormant ETO requires a larger rent payment.

Proposition 1. *The rent payments ($\widetilde{R}$) associated with a lease that has a strictly positive probability of ETO exercise are higher than the rents under a lease without an ETO (R). Hence,*

$$\widetilde{R} - R > 0 \quad \text{with } \widetilde{R} = R\left[1 + \frac{\sum_{s=t+1}^{T} p(t,s)\mathbb{Q}(\tau \leq s)\mathbb{Q}(s \leq \eta)}{\sum_{s=t+1}^{T} p(t,s)\left[\mathbb{Q}(s < \tau) + \mathbb{Q}(\eta < s)\right]}\right] \tag{3.4}$$

where $p(t,s) = \mathbb{E}_{\mathbb{Q}}\left(\frac{1}{B(s)}\right)B(t)$, and ETO notification and re-leasing follow Poisson processes with respective intensities λ_τ and λ_η under $\mathbb{Q}$.

Proof. See Appendix B.1 for the derivations. □

[11] The baseline environment assumes no grace period, $\alpha = 0$, and that the landlord faces no lapse in re-leasing if the ETO is perceived as a zero-probability event. We refer to this as a dormant ETO. In addition, the baseline model treats the exercise and notification of the ETO equivalently.

3.3 Property Value Loss

Recognizing that the current property value reflects the risk-adjusted sum of discounted cash flows and the liquidation value, we formalize the loss in property value arising from the underpricing of government contract risk associated with the ETO. Under the assumption of independence among the ETO notification, re-leasing, and the zero-coupon bond price processes, we derive a closed-form expression for the loss in property value.

Proposition 2. *Suppose the time t value of a property without ETO pricing is $V(t)$, and with ETO pricing the value is $\widetilde{V}(t)$. Denote the time t discounted value loss induced by mispricing the ETO as $L(t) = \frac{V(t)-\widetilde{V}(t)}{B(t)}$. With ETO notification and re-leasing following Poisson processes having respective intensities λ_τ and λ_η under $\mathbb{Q}$, the closed-form solution of the loss is*

$$L(t) = R \sum_{s=t+1}^{T} p(t,s) \left[e^{-\lambda_\eta s} - e^{-(\lambda_\tau + \lambda_\eta)s} \right] \tag{3.5}$$

Proof. See Appendix B.2 for the derivations. □

3.4 CMBS Bond Price Adjustment

We consider a pool of N otherwise identical properties that differ only in the perceived probability (salience) of exercising an ETO. A fraction θ_0 of the properties are assumed to have a dormant ETO (i.e., a zero probability of exercise) while the remaining fraction $\theta_1 = 1 - \theta_0$ have a strictly positive probability of ETO exercise. Each dormant-ETO (respectively, ETO-salient) property generates rent R ($\widetilde{R}$) per period and a terminal liquidation value $V(T)$ ($\widetilde{V}(T)$) at maturity T, conditional on the lease remaining active.[12]

We assume that the ETO exercise time τ for positive-ETO properties follows a Poisson process with constant intensity λ_τ under the risk-neutral measure $\mathbb{Q}$. The survival probability of the lease to time s conditional on surviving up to t, is then:

$$\pi(s) = \mathbb{Q}(\tau > s \mid \tau > t) = e^{-\lambda_\tau (s-t)} \tag{3.6}$$

where $\pi(s) = 1$ for dormant-ETO properties. The aggregated cash flow from the pool at time $s > t$ and terminal date liquidation value are, respectively:

$$\Gamma(s) = N \left[\theta_0 \widetilde{R} + \theta_1 R \mathbb{1}_{\tau > s} \right] \tag{3.7}$$

[12] We abstract from other elements commonly featured in CMBS pricing models, such as prepayment and vacancy risks, default risk, and heterogeneous recovery rates, as a full treatment of CMBS bond pricing is beyond the scope of this paper. For comprehensive discussions of CMBS pricing frameworks incorporating these additional dimensions, see, for example, Christopoulos et al. (2008) and Diener et al. (2012).

$$\Psi(T) = N\left[\theta_0 \widetilde{V}(T) + \theta_1 V(T)\mathbb{1}_{\tau>T}\right] \tag{3.8}$$

The no-arbitrage price of the CMBS bond at time t, denoted $\phi(t)$, equals the risk-neutral expected present value of the aggregated future cash flows and terminal liquidation value:[13]

$$\phi(t) = \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{T} \frac{\Gamma(s)}{B(s)} + \frac{\Psi(T)}{B(T)}\right] \tag{3.9}$$

Substituting in for $\Gamma(s)$ and $\Psi(T)$ in equation (3.9), we can write

$$\phi(t) = N\left[\sum_{s=t+1}^{T} p(t,s)\left(\theta_0\widetilde{R} + \theta_1 R e^{-\lambda_\tau(s-t)}\right) + p(t,T)\left(\theta_0\widetilde{V}(T) + \theta_1 V(T)e^{-\lambda_\tau(T-t)}\right)\right] \tag{3.10}$$

The following proposition establishes that a marginal increase in the ETO notification arrival intensity (i.e., an increased likelihood of exercising ETOs) has a negative effect on the CMBS bond price.

Proposition 3. *Let $\phi(t)$ denote the arbitrage-free price of a CMBS bond backed by a pool of N properties, where a fraction θ_1 are subject to early termination option (ETO) risk with Poisson intensity λ_τ. Then $\phi(t)$ is strictly decreasing in λ_τ, that is,*

$$\frac{\partial\phi(t)}{\partial\lambda_\tau} < 0 \tag{3.11}$$

Proof. See Appendix B.3 for the derivations. □

3.5 Regime Switching

We extend our baseline model by allowing the ETO notification arrival intensity, λ_τ, to follow a two-state continuous-time Markov chain, capturing a low-intensity (L) and a high-intensity (H) federal lease cancellation regime. The low-intensity state reflects the baseline policy environment prevailing before the DOGE initiative, while the high-intensity state represents the aggressive exercise of ETOs under the DOGE-driven exit strategy. The regime process $X_t \in \{L, H\}$ has the transition probability matrix:

$$Q = \begin{pmatrix} -q_{LH} & q_{LH} \\ q_{HL} & -q_{HL} \end{pmatrix}, \quad \lambda_\tau(t) = \begin{cases} \lambda_\tau^L & X_t = L \\ \lambda_\tau^H & X_t = H \end{cases}$$

Economically, q_{LH} reflects the hazard of federal leasing policy becoming more aggressive in the *exit regime* (e.g., the beginning of the 2025 DOGE-led terminations), while q_{HL} reflects

[13] In our empirical setting, we focus on the first-loss tranche, as it is the most junior in the CMBS waterfall structure and absorbs cash flow losses first when the market becomes ETO-salient.

reversion to the *baseline regime* (e.g., the rescission of some DOGE terminations). Define $u(s,i) = \mathbb{Q}(\tau > s \,|\, X_t = i)$ as the lease survival probability at time s in regime i, which satisfies the system of ordinary differential equations,

$$\frac{\partial u(s,i)}{\partial s} = -\lambda_\tau^i \, u(s,i) + \sum_{j \neq i} q_{ij} \big(u(s,j) - u(s,i) \big)$$

The term $-\lambda_\tau^i \, u(s,i)$ represents hazard-driven decay of survival in regime i, while the term $\sum_{j \neq i} q_{ij} \big(u(s,j) - u(s,i) \big)$ accounts for regime switching, with transition between states at rates q_{ij}. Incorporating these extensions into (3.6), we obtain the probability of the lease surviving to time s conditional on surviving up to t and being in regime i:

$$\widehat{\pi}(s,i) = \mathbb{Q}(\tau > s \mid \tau > t, X_t = i) \tag{3.12}$$

Applying the state-dependent conditional survival probabilities directly to aggregate cash flows in (3.7) and the terminal liquidation value in (3.8), we derive the no-arbitrage price of the CMBS bond at time t, conditional on starting in regime i, denoted $\widehat{\phi}(t,i)$:

$$\widehat{\phi}(t,i) = N \left[\sum_{s=t+1}^{T} p(t,s) \big(\theta_0 \widetilde{R} + \theta_1 R \widehat{\pi}(s,i) \big) + p(t,T) \big(\theta_0 \widetilde{V}(T) + \theta_1 V(T) \widehat{\pi}(T,i) \big) \right] \tag{3.13}$$

The following proposition establishes that the CMBS bond price declines with a higher transition rate into the exit regime (i.e., aggressive ETO-canceling regime), but increases with a higher transition rate back to the baseline regime.

Proposition 4. *Let $\widehat{\phi}(t,i)$ denote the arbitrage-free price of a CMBS bond backed by a pool of N properties at time t in regime $i \in \{L, H\}$, where a fraction θ_1 are subject to ETO risk under a regime-switching intensity process with transition rates q_{LH} (from L to H) and q_{HL} (from H to L), and baseline and exit regimes' hazard levels be λ_τ^L and λ_τ^H respectively. Then, we obtain the following comparative statics:*

$$\frac{\partial \widehat{\phi}(t,i)}{\partial q_{LH}} < 0, \qquad \text{(higher likelihood of switching to the exit regime reduces bond price)}$$

$$\frac{\partial \widehat{\phi}(t,i)}{\partial q_{HL}} > 0, \qquad \text{(higher likelihood of returning to the baseline regime increases bond price).}$$

Proof. See Appendix B.4 for the derivations. □

3.6 Spatial Spillovers to Private Leases

We extend the regime-switching framework to incorporate a *spillover effect* of ETO notifications, whereby cash flows of nearby properties, including those without direct ETO exposure, are also adversely impacted. For tractability, we modify the original pool set up in Section 3.4 such that non-ETO properties (θ_0) earn safe cash flows $\bar{R}$ and valuation $\bar{V}(T)$ which are reduced proportionally by a contagion factor:

$$\Gamma_{\text{NonETO}}(s,i) = \theta_0 \bar{R} \varsigma \, \mathbb{1}_{\{\tau<s\}} \qquad \Psi_{\text{NonETO}}(T,i) = \theta_0 \bar{V}(T) \varsigma \, \mathbb{1}_{\{\tau<T\}}$$

where $\varsigma > 0$ denotes the proportional ETO-induced spillover loss rate experienced by non-ETO properties. The larger this term, the more aggressively the government is perceived to behave. The parameter ς measures the *sensitivity of nearby (non-ETO) properties* to the observed (delivered) ETO notifications. If $\varsigma = 0$, then non-ETO properties are fully insulated from the negative spillover effects induced by ETO notifications. If $\varsigma > 0$, then cash flows of non-ETO properties decline proportionally to both realized ETO activity $(1 - \pi(s,i))$ and the spillover rate ς. The adjustment factor $(1 - \varsigma(1 - \pi(s,i)))$, denoted as Ξ hereafter, thus scales down otherwise safe cash flows as the contagion effect grows.

In practice, negative spillover effects to income generated by nearby non-ETO properties can occur if there are consumption externalities due to the relocation of government employees who would have otherwise visited nearby businesses while traveling to and from work (Miyauchi et al., 2025). In Section 6.4, we find no evidence in support of this mechanism in our analysis of retail foot traffic data. However, there could also be positive production spillovers to local contractors from having GSA tenants nearby (Duranton and Kerr, 2018), in which case lease cancellations could limit agglomeration effects on the margin of new business entry, thus increasing vacancy rates for non-GSA properties in the same neighborhood.

We can thus derive the arbitrage-free contagion-adjusted price of the CMBS bond at time t, conditional on starting in regime i as

$$\widehat{\phi}_c(t,i) = N \left\{ \sum_{s=t+1}^{T} p(t,s) \left[\theta_0 \bar{R}\, \Xi(s,i) + \theta_1 R\, \widehat{\pi}(s,i)\right] + p(t.T) \left[\theta_0 \bar{V}(T)\, \Xi(T,i) + \theta_1 V(T)\, \widehat{\pi}(T,i)\right] \right\} \tag{3.14}$$

Proposition 5. *Let $\widehat{\phi}_c(t,i)$ denote the arbitrage-free price of a CMBS bond at time t, conditional on starting in regime i, in the presence of regime-switching ETO risk and spillover effects at rate ς. Then:*

1. *The bond price is strictly decreasing in the spillover parameter ς:*

$$\frac{\partial \widehat{\phi}_c(t,i)}{\partial \varsigma} < 0. \tag{3.15}$$

2. *The bond price is strictly increasing in the per-period adjustment factor $\Xi = 1 - \varsigma(1 - \pi(s,i))$ (i.e., the proportional share of unaffected non-ETO cash flows):*

$$\frac{\partial \widehat{\phi}_c(t,i)}{\partial \Xi} > 0. \tag{3.16}$$

Proof. See Appendix B.5 for the derivations. □

The second part of Proposition 5 motivates our use of a bond pool-level exposure share in empirically estimating spillover effects ς.

3.7 Testable Hypotheses

Building on the key results from the preceding sections, we formulate our three main empirically testable hypotheses.

Hypothesis 1 *A more aggressive federal lease cancellation policy, which increases the salience of early termination risk, leads to greater expected losses at the property level, which translates into lower CMBS bond prices.*

This testable prediction is directly motivated by Proposition 3 and Proposition 4. It implies that the CMBS bond price $\phi(t)$ declines as the early termination intensity λ_τ increases, since heightened termination risk reduces the expected stream of property-level cash flows. Consequently, an increase in λ_τ, driven by more aggressive DOGE policy actions, disproportionately erodes the value of properties with embedded ETOs through elevated expected losses. These losses propagate to the CMBS market by lowering the risk-adjusted present value of the pooled cash flows, thereby depressing bond prices.

Hypothesis 2 *Properties that receive ETO notifications exhibit a decline in net operating income (NOI), relative to otherwise comparable properties, reflecting the direct operational impact of government lease termination risk on property-level cash flows.*

This prediction is theoretically grounded in Proposition 1 and Proposition 2, which formalize the loss in property value arising from the mispricing of government contract risk embedded in the ETO. These results demonstrate that the present value of property-level cash flows declines when the risk of ETO exercise is underestimated, as reflected in higher expected losses due to premature lease terminations and uncertain re-leasing outcomes. Economically, this mechanism operates through a reduction in expected rental income streams and an increase in vacancy risk, both of which lower the landlord's net operating cash flows. Properties subject to ETO risk experience direct operational disruptions as leases terminate earlier than anticipated and re-leasing takes time, during which rent income is forgone. These property-level income shortfalls aggregate up to large declines in net operating income, property values, and ultimately CMBS bond prices, as we show via simulations in Section 7.

Hypothesis 3 *Properties located near ETO-notified buildings exhibit a decline in net operating income (NOI), property values, and associated CMBS bond prices, relative to otherwise comparable properties, reflecting the spillover impact of government lease termination risk on nearby, non-notified assets.*

The intuition behind this prediction is captured by Proposition 5, which formalizes the adverse spillover effects of ETO notifications on nearby, non-notified properties through a contagion mechanism. These results demonstrate that the present value of property-level cash flows and valuations for non-notified assets decline as the spillover intensity increases, reflecting market perceptions of heightened vacancy risk and reduced tenant demand in the surrounding area. Economically, this mechanism operates through localized deterioration in property market fundamentals, as nearby tenants reassess location desirability and avoid DOGE-notified areas altogether. Non-notified properties experience indirect income disruptions as occupancy and rental rates weaken, even if no direct lease termination occurs. We show in most simulations in Section 7 that the largest share of total value destruction originates from spillovers to ETO-eligible but retained leases rather than losses accruing to properties with DOGE-notified leases.

4 Commercial Property Market Data

This section describes the data and the construction of key measures used in our difference-in-differences regressions. We provide summary statistics for our estimation samples and document several stylized facts about the GSA-linked CMBS market.

4.1 Data Construction

Our sample consists of three main datasets: (i) government cancellation information from the DOGE website, (ii) bond, loan, and property information from Trepp, and (iii) federal lease inventory information from the GSA website. As of writing, our data cover the months up to and including June 2025.

DOGE dates. Since March 23, 2025, we have hand collected twice per day (once in the morning and once in the evening) the DOGE website's information listed under its real estate savings section. The website provides the main agency, location, square footage under lease contract, and an estimated total savings amount involved in an early cancellation or lease non-renewal. Crucially, we collect the notification-sent date of the early termination option that the government plans to exercise, which is the basis for our treatment definition in the regression analysis. We use various media sources, plus website snapshots from the Wayback Machine and the JLL Federal Lease Termination Tracker, to backfill date announcements and lease information since the initial publication of savings on DOGE's website on February 18, 2025.

CMBS price data. We use the Trepp CMBS dataset which offers month-end bond values at the CUSIP level. We focus on monthly data from Trepp rather than higher frequency data on CMBS spot markets from the Financial Industry Regulatory Authority's Trade Reporting and Compliance Engine (TRACE), since the panel is completely unbalanced at daily frequency due to lack of liquidity in CRE debt markets. Restricting to a balanced panel is important for the interpretation of our results to the extent that bond pools may be restructured in response to DOGE announcements, meaning that we may not be able to plausibly separate mortgage composition effects from the role played by the activation of the dormant ETO clause in the repricing of the CMBS market. Nonetheless, in the Appendix we implement daily frequency event studies using the daily trade summary data from TRACE for those CUSIPs represented in Trepp. In doing so, we still find that bond prices for CUSIPs directly implicated in canceled leases fall after DOGE termination actions.[14]

Trepp FEED data. To construct a comprehensive panel dataset of government real estate investments and property performance, we supplement the CMBS bond pricing data from Trepp with the set of four tables contained in their FEED data: the *bond*, *deal*, *loan*, and *property* tables.

We merge across tables in Trepp according to the following steps:

1. We match CUSIPs in the *bond* table to the corresponding panel of CUSIPs in the Trepp CMBS file to obtain a bond-price table. This step effectively filters our sample to include only loans securitized into bonds with transaction prices.

2. We then many-to-one join the bond-price table to the *deal* table based on the deal name and tape date to pick up information on the performance of the loans within each CMBS bond pool. Each deal can result in multiple bonds.

3. We merge the *property* and *loan* tables in a many-to-many fashion using the Trepp deal name, the unique loan identifier, and the distribution date. The merge is many-to-many because one property can have multiple loans used to finance it, and, in turn, the same loan can be used to finance several different properties.

4. Finally, we create a bond-loan-property table by many-to-many matching the property-loan table with the bond-deal table created in step 2, thus keeping only properties and loans corresponding to a bond-deal. We perform this match using the Trepp deal name contained in both sub-tables and by matching the loan distribution date to the bond tape date.[15]

Merging across all five Trepp tables allows us to recover measures of performance, such as net operating income (NOI) or net cash flow (NCF), at both the loan level and averaged across properties within the same bond deal, and at different points in time (e.g., as of securitization vs. as of the last CMBS reporting date) to measure lease characteristics as close to DOGE announcements as possible. We drop from the sample the roughly one-fifth of loans where the NOI is only reported as of the securitization date, since for such loans we cannot isolate changes in cash flow performance. We winsorize bond prices at the 1st and 99th percentiles throughout our analysis. Similarly, to account for skewness in the distributions, we winsorize cash flow-related variables, such as NOI and debt service coverage ratios (DSCR) at the median plus or minus five times the variable's interquartile range.

CMBS bond ratings. We follow the classification scheme suggested by Flynn and Ghent (2018) to divide bonds into three seniority categories (first-loss group, mezzanine, and senior). We define

[14] The TRACE data are typically released with a three to four-month lag. Given the real-time nature of our exercise, this is another reason why we use the monthly Trepp CMBS data to form our main bond price panel.

[15] The loan distribution date and bond tape date match in 100% of cases in Trepp.

the first-loss group (FLG) as consisting of tranches which have a rating of CCC, or CCC+, or those which are unrated. In cases where the bond receives multiple agency ratings, we use the S&P rating. If the S&P rating is unavailable, we use the Fitch rating. If both the S&P and Fitch ratings are unavailable, we adopt the Moody's rating.[16] We refrain from using Kroll and Morningstar DBRS ratings in our baseline setup given 2023 SEC charges against these agencies regarding record-keeping failures for CMBS transactions.[17]

GSA leases. The U.S. General Services Administration (GSA) publishes its lease inventory information on a monthly basis. GSA lease inventory consists both of space leased by federal agencies as well as by private contractors engaged in government business. We compile a panel of GSA leases from January 2015 to the present by appending all the monthly inventory files. We collapse the GSA lease inventory panel to the unique geolocation level and separately record lease effective, expiration, and DOGE termination dates for individual leases within the same property. We then many-to-one match on the standardized property address the collapsed GSA inventory panel to the bond-loan-property table for use in our loan-level analysis.

To link GSA leases with properties in the Trepp CMBS dataset, we generate candidate matches using multiple criteria: (i) string similarity between standardized GSA and Trepp addresses, (ii) geospatial proximity between GSA and Trepp coordinates (the latter obtained from the Google Maps API), (iii) exact matches in square footage for the five largest tenants, and (iv) exact matches in lease expiration dates for these same tenants. We manually review approximately 400 potential matches for the Washington, D.C. market to validate correct pairings.[18] This procedure yields 91 matched leases (seven terminated) for D.C. and 45 leases (seven terminated) for Atlanta, corresponding to average match rates of roughly 25%. We believe the lower match rate in D.C. is primarily due to the fact that Trepp only covers CMBS-securitized properties, rather than match quality, given the broad criteria used to define the initial match pool.

In Table 1, we list the federal agencies who are tenants in DOGE-notified and ETO-eligible leases by cross-referencing their Trepp property address with online listings from CoStar. The agencies with office space within DOGE-terminated leases contained in a securitized pool include, on top of GSA divisions, Housing and Urban Development (HUD), Veteran's Affairs, Department of Homeland Security, Federal Emergency Management Agency (FEMA), Department of the Treasury, Internal Revenue Service (IRS), Department of Energy, Federal Energy Regulatory Commission, and Federal Aviation Administration (FAA) offices. Leases in our estimation sample for Washington, D.C. span 36 distinct federal agencies.

[16]We obtain similar tranche classifications – and therefore quantitatively similar results – if we instead divide up bonds into tranches based on the consensus grouping across rating agencies. The tranche group classifications overlap in 86% of cases between S&P and Fitch, 89% of the time between S&P and Moody's, and in 76% of cases between Fitch and Moody's.

[17]See the official SEC press release here: `https://www.sec.gov/newsroom/press-releases/2023-211`. That said, our results are similar and our sample size is larger if we include the Kroll ratings to classify bonds.

[18]Note that property geo-coordinates provided by Trepp pertain to the centroid of a zip code, not an address. Hence, we use Google Maps API to obtain more precise locations.

In Figure 12 we map the spatial distribution of canceled leases relative to the overall CMBS-linked property market in Washington, D.C. The seven canceled leases are clustered in the downtown areas of the city around the National Mall.[19] These seven terminated leases are attached to 74 unique CMBS CUSIPs. On top of the seven leases which actually received a letter from DOGE, ETO-eligible leases in the D.C. area span 285 CUSIPs across 15 bond pools on the eve of DOGE's formation, pointing to the potentially broader impact of DOGE's announced terminations despite the small number of leases involved. In what follows, we will distinguish currently ETO-eligible leases from soon-to-be ETO-eligible leases. We define soon-to-be-eligible leases those which will enter the soft term of the lease at any point during the current presidential administration, or between January 2025 and January 2029. We indicate "close" properties within a 1-mile radius of the affected leases and relatively "far" properties in the metro area but still within a 5-mile radius. These rings form the basis for our implementation in Section 5.3 of a difference-in-differences strategy to isolate spatial spillovers of the DOGE actions.

CoreLogic Tax data. We use the property tax roll from CoreLogic to assess the quality of our merge between GSA leases and Trepp properties. CoreLogic Tax (recently renamed "Property Basic") consists of a panel of all parcels assessed for property taxes, including those which are ultimately determined to be tax-exempt. We keyword parse the use descriptions in CoreLogic to identify office properties. We then compare within each metro area the fraction of GSA properties matched to Trepp records with the ratio of Trepp properties to CoreLogic office properties. This yields a measure of the representation of GSA relative to non-GSA leased properties in CMBS submarkets.

SafeGraph/Advan foot traffic. In our analysis of the spillover effects of DOGE cancellations to neighboring properties, we obtain data from Advan Weekly Patterns (Advan Research, 2022), formerly SafeGraph, to measure foot traffic to nearby businesses. Advan offers foot traffic for points of interest (POI) such as retail chains and local businesses and amenities based on anonymized geolocated cell phone "pings."[20] We first identify POIs located within a radius of terminated GSA leases. We then isolate the weekly number of visits to each POI within the radius, restricting to a balanced panel of businesses with strictly positive weekly foot traffic in the 24 months prior to DOGE's formation. To isolate retail POIs and test for consumption externalities, we use the 2-digit NAICS codes corresponding to *Retail Trade*, *Arts, Entertainment, and Recreation*, and *Accommodation and Food Services*.

[19] Two sets of canceled leases are located in the same building complex, meaning there are five unique properties implicated.

[20] See recent applications of the SafeGraph data to study agglomeration spillovers from grocery store openings (Qian et al., 2024) and disparities in bank branch access (Sakong and Zentefis, 2025). Hou et al. (2025) discuss advantages and best practices for using this dataset in corporate finance applications.

4.2 Sample Construction

Our main estimation sample includes only those observations for which we are able to merge underlying properties from the Trepp dataset to entries in the GSA inventory. Individual properties may have multiple lease agreements – up to seven in the Washington, D.C. sample. We keep track of the lease sequence and associated dates for each tenant tied to the property. We include all leases active at any point from 2020 to 2025, which captures both active and expired agreements. Given that cancellation rates in the pre-DOGE period were relatively low (see Figure 9), this means there are often multiple GSA leases located at the same property.

We record several key dates related to each lease. The DOGE announcement date captures the date a lease termination was first published on the DOGE website. We record any dates that a termination was formally rescinded. We also collect the date DOGE reportedly sent each termination notice, which typically precedes the public announcement on the website. This is the key determinant of treatment status in our analysis, as a formal notification commences the government's intent to exercise the early termination option, which influences CMBS bond prices and property net operating income immediately.[21] These variables are supplemented by GSA inventory dates, including the lease effective, lease expiration, and termination right dates, which are available for all leases regardless of termination status.

We successfully link 91 out of the 286 GSA leases in the 2020–2025 inventory for Washington, D.C. to Trepp properties, yielding a match rate of 31.8%. As a point of reference, Trepp's D.C. commercial property coverage in 2023 accounts for only about 6% of CoreLogic-parcel-identified CRE (or 3.4% if apartments are included). These tabulations suggest that GSA properties are overrepresented in the CMBS universe relative to non-GSA leased properties, especially in Washington, D.C. In our baseline analysis, to maximize statistical power, we use an unbalanced panel of bonds and properties, meaning that some bonds do not have valuations in each month. Our results are virtually identical if we restrict to a balanced panel, as nearly all the CUSIPs tied to the D.C. area have monthly Trepp valuations.

4.3 Summary Statistics

Table 2 reports nationwide descriptive statistics for the log of CMBS bond prices and the log of net operating income, stratified by ETO (early termination option) notification status and tranche group. Panel (A) presents the full sample, showing that the average log bond price is 3.74 with a standard deviation of 1.62 across approximately 17.7 million observations. The average log NOI is 15.04 with a standard deviation of 1.98, based on roughly 1.54 million observations.

Panel (B) focuses on leases where the ETO is exercisable but no notification is sent. Bond prices and NOI vary substantially across tranche groups. The mezzanine tranche has the highest

[21] See Figure 3 for examples of ETO and immediate rent abatement provisions in GSA leases.

average log bond price (4.09), while the first-loss and senior tranches report similar means, 3.28 and 3.25, respectively. The NOI is highest for the mezzanine group (16.70) and lowest for the unclassified group (14.51). In contrast, Panel (C) displays the subsample in which ETO notification was sent. Here, bond prices are lower overall for the first-loss group (2.98) and mezzanine group (3.92), while the senior and unclassified groups show slightly higher averages. Interestingly, NOI is more stable across tranche types in Panel (C), with a notable reduction in variance, especially for the first-loss group, standard deviation of only 0.15. The data indicate that CMBS bond pricing and property income levels vary meaningfully by risk tranche and government action. Notably, properties in the first-loss group exhibit significantly lower bond prices and more compressed NOI variation following ETO notification, consistent with market reassessment of risk upon government exit signals.

We focus our empirical analysis on the Washington, D.C. metro area, which contains about one-third of the GSA-leased portfolio. There are two primary motivations for this focus. First, as illustrated in Figure 5, Washington, D.C. ranks first among the top 20 states in terms of average square footage subject to early termination notifications. Second, Figure 7 reports the top 20 states by average single-tenancy concentration, with Washington, D.C. ranking 20th. Taken together, these figures suggest that the Washington, D.C. market presents a compelling case for initial study: it combines both high exposure to ETO risk and a relatively high degree of single tenancy at the building level, conditions that make it particularly salient to investors concerned about risks of portfolio exposure to government involvement in the real estate market.

For D.C. agency offices, DOGE communications occurred in tightly clustered periods and predominantly involved leases with substantial floor areas. All DOGE website announcement dates occurred on February 16, when the website first went live. Notification dates, or when DOGE sent letters to landlords indicating cancellation decisions, occur between January 30 and March 4, or between January 30 and February 13 when restricting to securitized properties. Average rented areas for leases range from 271,754 to 722,871 square feet across the notification date waves within this two-week time span.

The consistent timing of announcements and narrow date ranges for notifications underscore the centralized and discrete nature of the government's termination actions. This temporal clustering, combined with clear variation in treatment exposure (i.e., whether or not a lease received official notification), supports the use of a difference-in-differences design in our study. Our design exploits sharp and plausibly exogenous timing of treatment, allowing us to compare outcomes between terminated and non-terminated leases before and after the government's official announcement to remove agencies from private office spaces. The uniformity of treatment dates within the same month-long time period also mitigates concerns about staggered timing bias, strengthening identification of the causal effects.

5 Empirical Strategies

We adopt several difference-in-differences (DiD) research designs to test the two main hypotheses of the model outlined in Section 3.7. We define the treatment group as those leases for which the government officially issued an early termination option (ETO) notification, while the control group consists of otherwise similar leases which are soon-to-be past their termination right date and thus did not receive such a notification. The core identification logic rests on the premise that the issuance of an official ETO notice serves as a salient signal of the government's intent to invoke contract flexibility, thereby prompting an immediate market reassessment of risk.

To ensure valid comparisons across treatment and control groups, we focus on GSA leases with an exercisable ETO and those associated with first-loss group tranches of CMBS bonds in our baseline analysis. This sampling strategy is motivated by the assumption that the salience of government contract risk is most pronounced in settings where leases retain the legal flexibility for early termination and where the corresponding bond tranches are most sensitive to fluctuations in net operating income (NOI). In this context, soon-to-be ETO-eligible leases serve as a credible locus for identifying the "wake-up call" effect, wherein the market re-evaluates risk exposure upon recognizing the operational relevance of ETO provisions.

To isolate spillovers to the other groups of leases, such as already ETO-eligible leases and non-GSA leases, we use not-yet ETO-eligible leases as a common control group. Our choice of control group follows the logic that the latter set of leases are tied to debt contracts among the least exposed to the DOGE announcement shock. We validate this assumption by conducting contamination bias tests proposed by Goldsmith-Pinkham et al. (2024), and find no evidence of statistical cross-contamination of effects on bond prices or cash flows in comparing the directly notified and spillover groups to the not-yet eligible group of leases.[22]

We implement a spatial triple difference-in-differences (DDD) design to test whether the negative impact of government lease cancellations on CMBS valuations is disproportionately concentrated among properties in immediate proximity to the ETO-notified leases. The objective of this exercise is to test whether repricing of government contract risk is driven by a market-wide increase in salience, particularly through negative spillover effects on non-GSA leases located near ETO-impacted properties (**Hypothesis 3**). We estimate the average effect of government lease cancellations on CMBS bond prices for non-GSA leases located within 5 miles of a canceled lease relative to federal leases in the same region. Such a design captures the overall spatial spillover effect of government exit risk on private tenants.

[22]Our results are also quantitatively similar if we refine the definition of not-yet ETO-eligible leases to exclude leases which enter the soft term during the post-DOGE sample period.

5.1 Pooled Difference-in-Differences

Our main specification is a pooled difference-in-differences (DiD) regression of the form:

$$Y_{i,c,t} = \beta \cdot DOGE_{i,c} \times Post_t + \gamma \cdot Post_t + \eta \cdot DOGE_{i,c} + \xi' \cdot \mathbf{X_{i,c}} + \delta_{i,y} + \varepsilon_{i,c,t} \tag{5.1}$$

In this equation, $Y_{i,c,t}$ denotes the log of the bond price, net operating income, or debt service coverage ratio (DSCR) associated with deal i, bond CUSIP c, and time t.[23] The variable $DOGE_{i,c}$ is an indicator equal to one if the underlying lease is notified by DOGE for early termination, as indicated by the DOGE website and corroborated by other industry sources monitoring the website in real time. $Post_t$ equals one in periods following the initial wave of DOGE announcements in February 2025 or later.

The interaction term $DOGE_{i,c} \times Post_t$ is our key variable of interest. The coefficient β on the interaction term captures the average treatment effect of the DOGE notification on bond prices (or property performance) for tranches that are ETO-eligible relative to those that are not. The vector of pre-DOGE characteristics $\mathbf{X_{i,c}}$ controls for pre-existing differences in bond prices between DOGE-notified vs. non-notified tranches, such as initial differences in square footage, delinquency rates, and mortgage features.[24] In our most stringent specifications, we include 5-digit property zip code and deal closing year or bond CUSIP fixed effects, which absorb these control variables and the $Post_t$ dummy.[25] The vector of deal closing year dummies $\delta_{i,y}$ strips out any temporal variation in past CRE debt market conditions that might drive differences in the *ex post* performance of CMBS bonds.

Because our underlying data used to estimate (5.1) are at the bond-deal-property level, we experiment with clustering the standard errors at various levels. Standard practice would be to cluster at the level of treatment (Bertrand et al., 2004), which in this case is a lease contract. Yet, leases are not a unit of analysis in Trepp; individual leases can only be inferred based on tenants reported as occupying a property in a given quarter. We therefore ultimately cluster the standard errors at the bond CUSIP level, as doing so yields more conservative confidence intervals and accounts for the fact that the same bond can be tied to multiple leases in the dataset to the extent that the loan pools are not fully geographically diversified.[26]

[23]Note that the size-weighted average DSCR across properties within the bond pool is always strictly positive in our sample, so the log transformation does not result in us dropping any observations.

[24]DOGE-terminated leases tend to occupy large spaces and have higher debt service coverage for the attached mortgage, than their non-terminated but ETO-eligible counterparts, but higher loan-to-value ratios, 30-day loan delinquency rates, and longer delinquency spells conditional on missing a loan payment.

[25]Given the small number of securitized CRE properties relative to residential properties, including finer geographic fixed effects leads to many singleton cells. Therefore, we adopt 5-digit property zip code fixed effects as our baseline set of neighborhood fixed effects.

[26]For our cash flow-based outcomes of interest, we verify that our results quantitatively hold even if we collapse the panel down to the property level and estimate versions of (5.1).

We highlight that in equation (5.1) we consider treatment in this baseline specification to be both absorbing and to have common timing across all CMBS bonds. Hence, $DOGE_{i,c}$ is not indexed by t, and $Post_t$ is not indexed by i. We justify this simplifying assumption on the grounds that for the Washington, D.C. sample where many of the nationwide set of terminated leases are concentrated, DOGE sent all cancellation notices to landlords and tenants within a one-month timespan between January 30 and February 13, 2025. Given that our main data source covering CMBS prices and property fundamentals is at the monthly frequency, this effectively means treatment occurs within the same time period for the directly affected units.[27]

An alternative notion of treatment timing would use the date that terminated leases were listed under the “savings" section of the DOGE website. For many leases notified in the first wave of letters sent on January 30, the lease was not listed on the DOGE website until late March. Given that information about terminated leases would have already become common knowledge among investors in the intervening period – especially given extensive media coverage of the topic starting in early February (e.g., Associated Press, 2025; CoStar News, 2025a) – we consider the original information provided to the market on January 30 as the realized shock to the dormant ETO clause. To the extent that our main analysis uses monthly frequency data, using the DOGE website initiation date of February 18 as the timing cutoff would generate identical results.

5.2 Parallel Trends and Non-Anticipation

Our identification strategy relies on the parallel trends assumption that CMBS prices and loan and property performance metrics would have evolved similarly across the notified and non-notified groups of leases if not for the DOGE cancellations. We strengthen the credibility of this assumption by restricting the sample used to estimate (5.1) to leases that are subject to ETO provisions and by focusing on the first-loss group (FLG) tranches which are the most sensitive to property-level cash flow fluctuations. By ensuring that both groups are comparably at risk and similarly structured prior to the intervention, we mitigate concerns about underlying selection bias. Any divergence in outcomes following the notification date can then be interpreted as the causal effect β of the government's issuance of formal ETO notices.

We isolate the FLG tranches by excluding mezzanine and senior tranches in our estimation sample; the latter are less exposed to losses in the event a commercial mortgage defaults, a potential outcome in cases where delinquency is induced by prolonged vacancy after a sudden lease cancellation. Further, Ashcraft et al. (2019) show that buyers of B-pieces (low-rated CMBS tranches) act as gate-keepers in the CMBS market since they re-underwrite all the loans in the underlying pool, and for this reason B-pieces are more informationally sensitive. Restricting to

[27] Future work will examine whether the market reacts in a symmetric fashion to leases being removed from the DOGE website, using staggered treatment estimators à la de Chaisemartin and D'Haultfoeuille (2020) to implement versions of (5.1) in which $DOGE_{i,c,t}$ can reset from 1 to 0 when DOGE removes the lease from their list. Based on CRE listing company records, many leases originally characterized by DOGE as terminated but subsequently removed from the DOGE website are now being advertised as vacant.

the FLG allows us to zoom in on a segment of the CMBS market that would be the most likely to react to the increased salience of the ETO clause embedded in GSA leases.

To assess the validity of the parallel trends assumption after making our sample restrictions, we estimate an event study version of or baseline specification (5.1) that includes leads and lags of treatment and plot the resulting event-time coefficients β_t:

$$Y_{i,c,t} = \sum_{t=-4,t\neq-1}^{+3} \beta_t \cdot DOGE_{i,c} + \mu_{(i,c)} + \delta_{i,y} + \varepsilon_{i,c,t} \tag{5.2}$$

where $\mu_{(i,c)}$ refers to a set of fixed effects capturing different dimensions of the deal-bond combination, such as 5-digit property zip code, census block group, or CUSIP fixed effects. We set $t = -1$ to be the reference period, corresponding to January 2025. This reference period choice reflects the fact that the first set of DOGE terminations for the Washington, D.C. area was sent to landlords and tenants on January 30, 2025. Hence, the first month in which the market could react to the notifications is February 2025 ($t = 0$). Previewing the results, with or without our most stringent set of fixed effects, Figures 16 and 17 show no evidence of pre-treatment divergence for either bond prices or NOI. The same is true for DSCR in Figure 18. We cluster standard errors in each regression according to the unit at which the outcome is measured. For instance, clustering by bond CUSIP in regressions with bond prices as the outcome accounts for the fact that bonds are observed repeatedly over time and may exhibit persistent idiosyncratic factors such as liquidity, coupon terms, investor base, and trading frequency.[28]

A second identifying assumption is non-anticipation, which posits that market participants did not revise their expectations about lease termination risk prior to the observed announcement or notification dates. In our setting, this assumption is justified by the temporal clustering of ETO notifications and the widespread perception prior to the event that ETO clauses were legally available but operationally dormant. The low historical volatility and levels of early lease cancellations depicted in Figure 10 bolster our interpretation of lack of ETO salience in the pre-DOGE period. Hence, any observed bond price responses are likely triggered by the government's formal signaling of intent, rather than by market speculation or information leakage. The absence of pre-trends in average bond price and NOI trajectories across tranches in Figures 13, 14, and 15 further supports the view that the treatment was both unexpected and exogenous, thereby validating the use of post-notification variation as a natural experiment.

5.3 Spatial Difference-in-Differences

To account for the possibility of spillovers across space and CRE debt markets, we estimate a spatial difference-in-differences regression, distinguishing between GSA and non-GSA lease

[28]The pre-period coefficients are insignificant regardless of the level of clustering, including clustering by property zip code, deal id, or CUSIP.

performance:

$$Y_{i,c,r,t} = \beta \cdot Spillover_{i,c,r} \times Post_t + \gamma \cdot Post_t + \eta \cdot Spillover_{i,c,r} + \mu_{(i,c)} + \chi_{r,t} + \varepsilon_{i,c,r,t} \tag{5.3}$$

The dependent variables, subscripts, and $Post_t$ follow the same definitions as described for equation (5.1). The treatment variable in the spatial difference-in-differences is $Spillover_{i,c,r}$, an indicator equal to one if two criteria are satisfied: (i) the bond-deal involves a private tenant, and (ii) the underlying space being rented is located within a 5-mile radius of an ETO-exercised lease. The interaction term $Spillover_{i,c,r} \times Post_t$ captures the average spillover effect of DOGE lease cancellations on bonds not linked to GSA-leased properties relative to more distant units within the 5-mile radii of the ETO-exercised leases in the D.C. metro area. The index r emphasizes the fact that observations are connected to a particular ring r with the location of a DOGE-canceled tenant as its centroid.

As shown in the map of Figure 12, a property can be contained within multiple rings. We account for neighborhood-specific time trends that may overlap with the geography-based definition of treatment by including ring-by-time fixed effects $\chi_{r,t}$ in each regression. While our baseline analysis studying the direct effects on terminated lease properties relies on the Trepp data, in estimating (5.3) we include rings defined by non-securitized leases which were also canceled by DOGE (see the map in Figure C.3).

We further decompose the $Spillover$ dummy in (5.3) via a spatial triple difference-in-differences (DDD) design to determine whether there is any incremental impact of proximity to canceled leases within the innermost 1-mile radius relative to those in the surrounding 1-to-5 mile band. Equation 5.3 conflates two notions of treated status: one based on spatial proximity, and another based on whether the lease involves a private tenant. A natural question arises of which spillover margin is quantitatively more prominent. We answer this question by estimating the following equation:

$$\begin{aligned} Y_{i,c,r,t} &= \beta \cdot (Ring_{i,c,r} \times Private_{i,c} \times Post_t) + \gamma_1 \cdot Post_t + \gamma_2 \cdot Private_{i,c} + \gamma_3 \cdot Ring_{i,c,r} \\ &+ \gamma_4 \cdot (Private_{i,c} \times Post_t) + \gamma_5 \cdot (Ring_{i,c,r} \times Post_t) + \gamma_6 \cdot (Ring_{i,c,r} \times Private_{i,c}) \\ &+ \mu_{(i,c)} + \chi_{r,t} + \varepsilon_{i,c,r,t} \end{aligned} \tag{5.4}$$

where $Ring_{i,c,r}$ is an indicator equal to one if the property lies within one-mile radius of the ETO-exercised D.C. properties, and $Private_{i,c}$ indicates that the deal involves a non-GSA lease. The new coefficient of interest, β, captures the differential post-treatment effect for private bonds located within the treatment radius (1 mile), relative to other private bonds located further away (between 1 and 5 miles).

6 Main Results

We now present our main empirical results using the notification of DOGE cancellations of GSA leases as a natural experiment to test the hypotheses implied by our arbitrage pricing framework.

6.1 Results on Bond Pricing

Table 3 presents the main results from our difference-in-differences regression (5.1) examining the impact of early termination option (ETO) notifications on log CMBS bond prices. The coefficient on the interaction term, $DOGE \times Post$, captures the differential change in bond prices following the ETO notification for notified leases relative to not-yet eligible ones. Across columns (1) through (3), which sequentially include fixed effects for property zip code, deal-year and bond-month, tranche group, and bond CUSIP, the estimated treatment effect is negative and statistically significant at the 5% level. The estimated decline in bond prices ranges from 2.4 to 4.0 log points, suggesting that markets priced in heightened government contract risk immediately following the receipt of an ETO notice.

The relatively stable magnitude of the treatment effect across specifications, including the most saturated model, underscores the credibility of the causal interpretation. Including CUSIP fixed effects compares leases with vs. without a DOGE termination notice but bundled within the same tranche issue. Together, these results provide strong causal evidence that the government's ETO notification leads to a statistically and economically meaningful repricing of first-loss CMBS bonds, consistent with a market response to a previously mispriced contractual risk.

These empirical findings lend direct support to **Hypothesis 1**, confirming that an increase in government lease termination intensity λ_τ, realized through DOGE's ETO notifications, leads to a significant decline in CMBS bond prices. The observed repricing is consistent with the theoretical prediction that heightened policy-induced termination risk lowers the expected present value of property-level cash flows, which in turn diminishes the valuation of securitized tranches exposed to this risk. By documenting statistically and economically meaningful declines in first-loss tranche prices following ETO notices, the results validate the mechanism formalized in Proposition 3 and Proposition 4, demonstrating how mispriced contractual risk embedded in federal leases is revealed and incorporated into market prices through the DOGE policy shock.

We show in Appendix D that the repricing of government contract risk through DOGE's actions is also reflected in the prices and abnormal returns for REIT equities. We compare publicly-traded REITs holding D.C. office properties that are more or less exposed to the termination announcements based on the fraction of their NOI or leased square footage accounted for by GSA tenants. For the GSA-exposed REITs, cumulative abnormal returns (CARs) drop by 5% below normal levels, as measured in the months prior to the 2024 presidential election. The control group of D.C. office REITs with little to no GSA tenant exposure experiences a 2% increase in their CAR, reflecting that investors price the government contract risk rather than a general shift in fundamentals for the D.C. office market writ large.

6.2 Results on Property and Loan Performance

Table 4 reports results from estimating (5.1) with log net operating income (NOI) as the outcome. The analysis is again restricted to leases that are in the first-loss tranche group. The coefficient on the interaction term, $DOGE \times Post$, is consistently negative and statistically significant at the 1% level across all four specifications. The estimates imply that properties receiving ETO notifications experience a 5.1 to 6.8 log points decline in NOI, relative to comparable untreated properties, even after controlling for combinations of property zip code, deal, tranche, and bond-level fixed effects. The dynamic event study (5.2) shows a sharp drop in February 2025 ($t = 0$) NOI which has persisted up until May 2025, with a flat trend in cash flow prior to the creation of DOGE (Figure 17). These results indicate that government lease terminations exert significant downward pressure on building-level net income, validating the economic channel through which contract risk transmits into financial asset pricing.

Why is there an immediate negative effect on NOI of properties with canceled leases, despite the existence of a grace period granted to landlords subject to an ETO during which the government still pays rent? First, the grace period is negotiable, and some leases applied to agencies which were immediately disbanded (e.g., USAID). Second, as shown in Panel B of Figure 3, during the grace period the government is entitled to a rent abatement at either a pre-negotiated rate, or, in the case of net leases, at a rate based on a reduction in the operating expenses component of the rental payment in proportion to the leasable area being vacated. The typical rent abatement for a GSA lease subject to the vacant premises clause is between \$1.50 and \$2.00 per square foot (Holland & Knight, 2025). Under early 2025 estimates of market rent of \$60 per square foot for grade A office space in D.C., NOI would fall by roughly 3% due to the rent abatement, even without any immediate vacancy or changes to the property's operating costs. Hence, at most half of the initial drop in NOI we observe is due to rent abatement, and the other half is caused by agencies which left within a month of receiving a DOGE termination notice.

Declines in NOI for properties with terminated leases are mirrored in declines in debt service coverage ratios (DSCR), computed as a size share-weighted average across properties in the loan pool. We uncover a robust effect in Table 5 of a 2.5 to 2.7 log points drop in average DSCR for DOGE-notified bonds, with corresponding event study results in Figure 19. Similar to the event study for NOI, there is an immediate and persistent decline in DSCR for bonds exposed to a DOGE-notified lease.[29] Since creditors can call in their capital if DSCR falls below a threshold negotiated by the lender and any co-lenders in the intercreditor agreement, a deterioration of DSCR signals potential future covenant violations. This is an important concern due to the current

[29] We obtain qualitatively and quantitatively similar results if we instead use DSCR as a function of net cash flow (NCF) instead of NOI as the outcome variable.

high interest rate environment which has led to a tightening of interest coverage limits and constraints on firms' debt issuing capacity (Greenwald, 2019).

Since NOI is updated for most loans at a less than monthly frequency, our estimates for the effect of DOGE cancellations for this outcome and for DSCR – which is a function of NOI attached to loans included in the bond pool – are attenuated towards zero. For instance, in a typical month in the months leading up to the creation of DOGE, NOI experiences a month-on-month change for less than 10% of the loan sample. However, this fraction steadily rises in the post-DOGE period, peaking at 25% in May 2025. Based on lease effective dates and term lengths in the GSA lease inventory data, annual start dates for GSA leases active as of December 2024 are almost uniformly distributed across months in the year; leasing years are slightly less likely to start between February and March (when most DOGE announcements were concentrated) and more likely to start in August through October. Hence, the observed decline in NOI for terminated leases is due to DOGE actions rather than seasonality related to how frequently the loan reports cash flows for the underlying property.

Our empirical results are consistent with **Hypothesis 2**, supporting the prediction that underestimated government lease termination risk materially depresses property-level cash flows when exercised. The documented decline in NOI following ETO notifications aligns with the theoretical mechanism whereby premature lease terminations and heightened vacancy risk reduce expected rental income streams, impairing net operating performance. By showing that properties receiving ETO notices experience significant and economically meaningful income shortfalls relative to otherwise comparable properties, the findings validate the characterization in Proposition 1 and Proposition 2 in which mispriced contractual risk embedded in federal leases translates into operational disruptions and diminished cash flow realizations at the property level.

6.3 Spillover Effects to Nearby Properties

Table 6 presents the results of the spatial difference-in-differences regression estimating the impact of DOGE lease cancellations on CMBS bond prices. The analysis focuses on properties that are not directly leased to the GSA and lie within a 5-mile radius of the canceled leases in Washington, D.C. The interaction term, $Spillover \times Post$, exhibits a consistently negative and highly significant coefficient in the first two model specifications, with statistical significance at the 5% level. When we include more stringent fixed effects, the results are statistically significant at the 1% level. The estimates indicate that bond prices of first-loss tranches decline by between 9% to 11% compared to otherwise similar, unaffected tranches tied to private-tenant leases, even after controlling for property, ring-time, and bond-level fixed effects. The magnitudes of these bond price declines are larger than the direct impacts on DOGE-notified GSA leases, as estimated in the preceding subsection.

Table 7 documents the spatial triple difference-in-differences regression results from estimating the decomposed spillover equation (5.4). The analysis focuses on properties that are not directly leased to the GSA and lie within a 1-mile radius of the seven canceled leases in Washington, D.C.[30] As before, we subset the sample to properties located within a 5-mile radius buffer and the first-loss tranche group. As illustrated in Figure 12, this produces a set of treatment groups comprising leases located in the inner black rings and a control group situated between the inner black and outer purple rings.

The insignificant coefficient estimates on the triple interaction term, $Ring \times Private \times Post$, indicate that the price declines documented in the DiD estimates are not statistically stronger for properties located within the innermost 1-mile radius compared to those in the surrounding 1–5 mile band. In other words, while the baseline DiD results demonstrate economically and statistically significant price declines for private-sector properties in the broader treated area, the DDD estimates suggest that the negative impact is not stronger at closer distances. This finding is consistent with the interpretation that the spillover effects of federal lease cancellations are broadly felt within the 5-mile zone. Although we do not currently model intra-pool spillovers in our theoretical framework, the fact that the estimated negative coefficient on $Post \times Private$ drops by two-thirds with the inclusion of bond fixed effects in Table 7 suggests that spillovers are not driven by contagion across leases securitized into the same pool.

One plausible explanation for the insignificant triple-difference estimates is that the DOGE lease cancellation shock functions more as a market-wide increase in the salience of early termination options rather than as a hyperlocal shock to economic activity. Rather than affecting only the properties in the immediate 1-mile vicinity of a canceled lease, the DOGE initiative may have served as a systemic wake-up call to investors, triggering a broad reappraisal of federal lease risk across the Washington, D.C. commercial real estate market. The visibility and policy significance of the DOGE cancellations likely amplified investor attention to the ETO risk, even for properties that were not directly leased to the federal government. As a result, market participants may have repriced CMBS bonds across the entire 5-mile radius in a relatively uniform manner, reflecting heightened perceptions of risk, tighter underwriting expectations, or anticipations of reduced liquidity. In such an environment, the incremental informational content of physical proximity (e.g., being within 1 mile versus 3 miles) may be marginal compared to the broader behavioral and informational shock induced by the DOGE policy shift. This would imply that spatially granular variation in treatment intensity plays a secondary role relative to the general market reaction, thereby attenuating the differential effect captured by the DDD specification.

Our results in this subsection validate **Hypothesis 3** by showing that the sensitivity of CMBS bond prices to observed ETO notifications is empirically reflected in the price adjustments of

[30] Trepp classifies 95% of the properties in our spillover sample as being majority office use. Therefore, our analysis of spillovers within the CRE debt market is restricted to mostly office and mixed-use retail.

securities backed by private-tenant leases in the vicinity of ETO-exercised buildings. Our results underscore a pronounced spatial spillover effect, whereby government lease termination risk depresses the valuation of CMBS tranches collateralized by nearby, non-ETO properties. These findings are consistent with Proposition 5, which predicts that government lease cancellations trigger significant and economically meaningful declines in CMBS bond prices through the propagation of contract risk across space.

6.4 Testing for Consumption Externalities Using Retail Foot Traffic

The previous set of tests show that the distinction between relatively near vs. far non-GSA properties matters little for the magnitude of negative spillovers, pointing to a market-wide shock acting through debt markets. One possibility is that negative consumption externalities to nearby non-GSA leases are limited due to hybrid work modes put in place at terminated agency offices after the COVID-19 pandemic. Hybrid work policies are now widespread among private, publicly-listed firms (Flynn et al., 2024). If agency offices subject to DOGE lease termination feature hybrid work environments with a large fraction of hours worked from home, then there is little scope for spillovers to surrounding businesses arising from fewer employees making trips to and from the office.

To evaluate this conjecture, we classify federal agency tenants in Table 1 based on their return-to-office (RTO) policies listed in the August 2024 OMB Congressional Report (Office of Management and Budget, 2024). Both canceled and non-canceled but ETO-eligible federal tenants in the D.C. metro area are under RTO mandates stipulating that employees average between two and three days in the office each week. The limited cross-agency variation in telework mandates suggests that DOGE did not target leases for offices with substantially lower in-person attendance.

To directly test for the consumption externality channel, we implement a spatial DiD regression similar to (5.3) that compares the evolution of foot traffic for retail establishments located inside an inner ring to those located inside an outer ring. Specifically, for a given pair of radii $(\mathfrak{r}_{inner}, \mathfrak{r}_{outer})$ we estimate the following regression specification:

$$Y_{j,r,s,t} = \sum_{t=-10, t\neq -1}^{+10} \beta_t \cdot Spillover_{j,t} + \mu_j + \chi_{r,t} + \delta_{s,t} + \epsilon_{j,t,r,s} \tag{6.1}$$

In this equation, $Y_{j,r,s,t}$ denotes the volume of visits to POI j, located in outer ring r and belonging to 3-digit NAICS subsector s, during week t. The variable $Spillover_{j,t}$ is an indicator equal to one if the POI is located inside an inner ring (i.e., if the distance to a terminated DOGE lease is less than $\mathfrak{r}_{inner}$) and if the relative time to the termination of that DOGE lease equals t. μ_i and $\chi_{r,t}$ denote POI and ring-by-week fixed effects, respectively, and $\delta_{s,t}$ represents week-by-subsector fixed effects that absorb subsector-specific seasonality. We define the control group as those POIs that are located in an outer ring of radius $\mathfrak{r}_{outer}$ around a DOGE terminated lease by excluding POIs from the estimation that are located further than $\mathfrak{r}_{outer}$ away from a termination.

We account for the fact that our outcome of interest in (6.1) is a count variable using two alternative approaches. Our first approach uses a fixed-effects Poisson regression, as recommended by Cohn et al. (2022), and thus explicitly assumes that the outcome is sampled from a Poisson distribution. In our second approach we instead use OLS but with the log of weekly visits as the outcome, assuming that this transformed outcome is asymptotically normal distributed. For both approaches, we test for spillovers at different levels of proximity by varying the pair of inner and outer radii $(\mathfrak{r}_{inner}, \mathfrak{r}_{outer})$ from $(0.25, 0.5)$ miles to $(0.5, 1)$ miles.[31] Figure C.1 gives a visual presentation of the resulting rings. Throughout, we use Conley (2008) standard errors that are robust to the strong spatial autocorrelation present in the foot traffic data.[32]

Figure 20 displays a null effect on foot traffic after estimating (6.1) by either Poisson regression (Panel A) or OLS (Panel B), regardless of the choice of inner and outer radii. If anything, there is a slightly positive, but statistically insignificant trend in retail foot traffic in the weeks after DOGE announces a termination. We probe this null result further in Appendix C, but ultimately find that it holds even in a version of (6.1) in which we compare effects across early vs. late cohorts of DOGE terminations using the Callaway and Sant'Anna (2021) estimator (Figure C.4) or when we estimate ring DiD models for non-retail foot traffic (Figure C.2).

Overall, our analysis of RTO policies and foot traffic to nearby retail and non-retail establishments casts doubt on a consumption externality channel driving negative performance spillovers to non-GSA properties. Rather, the DOGE shock generated a revaluation of the office stock by activating a long-dormant contractual clause in leases.

7 Implied Property Value Losses

We analyze the property valuation consequences of government contract risk exposure using a simulation framework that is disciplined by the coefficient estimates from our main empirical results for the Washington, D.C. market to quantify the aggregate value losses implied by Proposition 2. Specifically, we focus on quantifying the key tail risks associated with ETO exposure by addressing two central questions: (i) what is the minimum loss incurred in the worst $\alpha\%$ of simulated outcomes, corresponding to the Value at Risk (VaR) at some percentile $1 - \alpha$, and (ii) what is the average loss conditional on being in this worst $\alpha\%$ of outcomes, known

[31] Precisely estimating spillovers at lower levels of proximity is not possible, as the effective number of observations in the small inner circle is insufficient due to spatial correlation. Estimating spillovers at higher levels of proximity is possible, but such estimates are likely contaminated by time-varying differences in foot traffic between the city center and the surrounding, less dense areas.

[32] In all estimations, we vary the maximal cutoff for spatial autocorrelation from 0.05 miles to 0.5 miles in 0.05 mile increments and report the standard errors that are the most conservative. Alternatively, clustering standard errors by ring-time or Census block group yields tighter confidence intervals across most specifications.

as the expected shortfall (ES).[33] Together, these risk measures allow us to assess the potential market-wide magnitude of adverse outcomes stemming from DOGE's lease termination actions.

We provide complete details for the simulation algorithm in Appendix E, but summarize the key steps here. First, we numerically simulate the loss distribution of $L(t)$ under stochastic ETO risk. Baseline property values $V(t)$, estimated using a hedonic pricing model, serve as counterfactuals in the absence of ETO repricing. Following the ETO regime-switching intensities described in Section 3.5, we model the evolution of property values over a five-year horizon using a jump-diffusion process. The Brownian motion component captures continuous market volatility, while the Poisson jump component with arrival intensity λ_τ and calibrated to observed DOGE notification rates represents abrupt lease terminations.

Each jump induces a proportional decline in asset value consistent with the empirically estimated 5–6% reduction in NOI for ETO-notified properties, as reported in Table 4. To account for spillover effects to the already ETO-eligible and non-GSA lease groups, we scale jump magnitudes by the empirically estimated declines in NOI for each of those groups (1.5% and 10.1%, respectively). Implicitly, in our simulation this means the group of not-yet ETO-eligible leases (the control group in our empirical setting) acts as a reference category. We conduct 50,000 Monte Carlo iterations, sampling with replacement from our estimation sample, to generate empirical loss distributions for both directly impacted and spillover assets. From these distributions, we compute the VaR and ES measures and compare losses to estimates of the overall value of office properties in the Washington, D.C. market and to implied taxpayer savings from canceled GSA lease payments.

7.1 Hedonic Property Valuation

As a preliminary step, we compute the lagged moving average of the four most recent appraised values relative to the securitization date for each property provided by Trepp:

$$\overline{V}_i = \frac{1}{4}\sum_{k=1}^{4} V_{i,t-k} \tag{7.1}$$

where $\overline{V}_i$ denotes the appraised value of property i at the k-th most recent observation. This average, $\overline{V}_i$, serves to mitigate idiosyncratic noise in any single observation and more closely approximates a steady state valuation for each property. Next, we estimate a standard cross-sectional hedonic pricing model:

$$\begin{aligned} \ln \overline{V}_i &= \beta_0 + \beta' \cdot \mathbf{X_i} + \varepsilon_i \\ \widehat{V}_i &= \exp\left(\widehat{\beta}_0 + \widehat{\beta}' \cdot \mathbf{X_i}\right) \end{aligned} \tag{7.2}$$

[33] This is also referred to as the conditional Value at Risk at the $1-\alpha\%$ level.

where $\overline{V}_i$ is the appraised value based on (7.1), $\mathbf{X_i}$ is a vector of property, lease, and location characteristics averaged over the preceding four quarters, and ε_i is an idiosyncratic error term. The estimated baseline value $\widehat{V}_i$ obtained from the regression fitted values captures both systematic pricing effects and property-specific deviations. We test the goodness-of-fit of our property value estimation via in-sample fit measures (e.g., adjusted R^2, RMSE, MAPE). Appendix E offers additional details on the hedonic regression and robustness analysis.

7.2 Modeling ETO-induced NOI Volatility

We simulate the loss distributions for three groups: ETO-eligible yet not exercised, ETO-exercised by DOGE, and non-GSA leased properties (“spillover"). These three groups capture the full spectrum of potential ETO-related losses: unrealized risk (ETO-eligible),[34] realized losses (DOGE-notified), and indirect spillovers (private-lease properties). Grouping them in this way enables us to quantify both the direct impact of DOGE’s lease terminations and the broader market consequences through spillover effects. Summed together, we obtain the total Value at Risk from the policy-driven disruptions to federal tenancy.

For properties that have been formally notified that the GSA has exercised its ETO, we model value trajectories with a jump-diffusion process that combines regular market volatility with the possibility of abrupt declines, capturing the sudden impact of federal lease cancellations as reflected in our empirical results in Section 6.2. For properties that are eligible for early termination but have not yet received a formal notification, we apply the same jump-diffusion framework but with an attenuated shock component, reflecting a heightened but less immediate exposure to termination risk. In the third scenario, we extend this framework to nearby private-tenant properties that are not directly subject to federal terminations but may nonetheless experience correlated declines through reduced demand, lower foot traffic, and negative signaling effects in local markets. In this case, shocks are dampened relative to directly affected assets, capturing the indirect yet meaningful disruptions that spill over into surrounding submarkets.[35]

To assess the medium-run impact, we capitalize the terminal property value using a discounted cash flow approach at a five-year horizon, treating the final simulated value as a stabilized income stream. Losses are then calculated as the difference between this impaired terminal value and a counterfactual benchmark based on the hedonic valuation in (7.2). We use a five-year horizon to match the typical length of the soft term for GSA leases. We simulate each of the three

[34] This group consists of federal leases that have termination right dates prior to the creation of DOGE but whose ETOs were not exercised by the DOGE.

[35] In a separate set of regressions, we estimate the coefficient estimates for the ETO-eligible and spillover groups and discipline our simulation parameters accordingly. In the first regression, the control group is defined as GSA leases that are ETO-eligible but not formally notified, with termination right dates falling before the creation of DOGE on January 20, 2025. We re-estimate our difference-in-differences spatial spillover regression with the aforementioned control group definition. See Appendix E for more details.

property scenarios 50,000 times using Monte Carlo methods to construct a distribution of possible outcomes. This approach allows us to estimate tail risks such as Value at Risk and expected shortfall under varying exposure intensities, thereby quantifying the systemic implications of federal lease termination risk across both government-leased and private real estate portfolios.

7.3 Simulation Results

We document that properties which are ETO-eligible or subject to spatial spillover effects are critical determinants of tail risk in property value declines. Figure 21 presents the simulated distribution of property value losses under early termination option (ETO) risk across three groups: already ETO-eligible properties (gray), ETO-notified properties (pink), and private-lease properties within a 5-mile radius subject to spatial spillovers (green). The estimated 95% Value at Risk (VaR) is \$256 million for ETO-notified properties, but much larger for the ETO-eligible sample at \$914 million, suggesting that formal notification exacerbates downside risk exposure. Private-lease properties exhibit a distinctly lower but still sizable 95% VaR of \$177 million, reflecting the vulnerability of assets indirectly exposed to ETO terminations via a negative shock to investor demand for holding CRE debt. The unified distribution underscores that while direct ETO exposure is consequential, secondary spillovers may pose even greater tail risks under systematic federal lease repricing conditions.

The simulation results reveal meaningful differences in the loss distributions, particularly with respect to the emergence of fatter left tails in the spillover group. These fat-tailed outcomes indicate elevated probability mass in the extreme loss region, suggesting non-linear amplification of risk among indirectly affected assets. Since the presence of such fat tails implies that risk cannot be fully captured by the first and second moments, it is important to investigate and quantify the expected shortfall that offers a more appropriate lens for stress testing a portfolio of ETO-exposed assets in a region.

7.4 Systemic Tail Risk and Market-Level Exposure

While the Value at Risk (VaR) provides a threshold for losses in extreme scenarios, a natural question is: what would be the expected magnitude of losses conditional on exceeding that threshold, that is, the average market-wide loss in the worst $1 - \alpha\%$ of simulated outcomes? To examine the average loss in nominal terms embedded in tail risks, we compute the corresponding expected shortfall (ES) in the worst 5% of outcomes, denoted as $\text{ES}_{0.95}(L)$, separately for each group of properties (i.e., ETO-notified, ETO-eligible, and Spillover) which is given by:

$$\text{ES}_{1-\alpha}(L) = \mathbb{E}\left[L \mid L \geq \text{VaR}_{1-\alpha}(L)\right] \tag{7.3}$$

According to (7.3) the 50%-level (median) expected shortfalls for ETO-eligible (non-exercised), ETO-exercised, and spillover properties are \$307 million, \$185 million, and \$84 million, respectively. The total expected shortfall across the three groups is approximately \$575 million. We report VaR and expected shortfalls for different levels of α% in Table 8. In tail-risk scenarios ($\alpha = 5\%$), five-year expected shortfalls total \$1.47 billion across the three groups of properties. Despite not being terminated, the ETO-eligible group comprises over half of the expected losses across the distribution of simulated portfolio draws. The reason is that ETO-eligible properties in D.C. enter into the DOGE period with lower occupancy rates, lower NOI, and more gross rentable area than their canceled counterparts, on average. The greater building scale and weaker starting fundamentals increase the scope for losses due to the repricing of government risk exposure.[36]

Placing this into the broader context of the Washington, D.C. metro area office real estate market, which spans roughly 158.6 million square feet and is valued at approximately \$40 billion based on price per square foot charged in recent transactions (BNP Paribas Real Estate, 2025; CommercialEdge, 2025; Cushman & Wakefield, 2025), this simulated loss represents a non-trivial 3.7% of total market capitalization in the worst 5% of scenarios.[37] These results highlight that early lease terminations are not merely localized or idiosyncratic events. Instead, they carry systemic implications through both direct value destruction and indirect spillovers across private-tenant assets.

Our average loss estimates are conservative for three main reasons. First, for each group of properties, our estimates of the average drop in NOI are attenuated towards zero due to loan cash flows in Trepp updating at a quarterly frequency for most loans. This empirical moment is the key parameter underlying the jump processes in the simulation. Second, our arbitrage pricing framework does not incorporate possible general equilibrium forces of lease cancellations. For instance, vacancies triggered by ETO-exercised buildings can initiate a cascading sequence of additional vacancies within the same building due to input-output networks (Duranton and Kerr, 2018) and/or hyperlocal consumption spillovers (Miyauchi et al., 2025). Such dynamics point to a structural vulnerability in regional CRE stability, driven by contagion effects induced via ETO exercises. Third, our reduced form empirical results and simulation analysis pertain to securitized properties. To the extent that negative spillovers can occur to non-CMBS properties which are typically smaller, the market-wide losses will be larger. Still, our results in Section 6.4 demonstrating no negative impact of DOGE cancellations on foot traffic suggest that the non-financial externalities of government contract risk are limited, at least in our setting.

Another way to contextualize the broader market exposure from ETO-induced value losses is to compare them against the magnitude of cost savings accruing to the federal government. At the peak of its lease termination campaign in March 2025, the Department of Government

[36] Similarly, even though the average decline in NOI for a private-tenant office property is several magnitudes greater than the decline for the other two groups, the typical privately-leased property is much smaller than a GSA-leased one, with a median appraised value of \$16.3 million compared to a median of \$156 million for the DOGE-notified subsample.

[37] The August 2025 CoStar Office Market Report for Washington, D.C. presents estimated total office asset values for the East End, CBD, and Tysons Corner areas of approximately \$49.8 billion.

Efficiency reported approximately $660 million in savings from federal lease cancellations and non-renewals nationwide (Politico, 2025). Notably, this figure is nearly matched by the $575 million in aggregate expected property value losses we simulate from *only the D.C. area* in the worst 50% of scenarios. Based on annual contract payments reported in the GSA lease inventory, at the height of lease terminations listed on the DOGE website as of mid-March 2025 the total savings based on the implied (non-discounted) annual lease payments remaining until lease expiration amount to $220 million, or $76.2 million when only including terminations in the broader D.C. metro area including Maryland, Virginia, and West Virginia. While the reported DOGE savings reflect meaningful fiscal relief on the public balance sheet, they come at the cost of potentially larger private-sector value erosion, concentrated in a regional office market already exposed to elevated federal tenancy risk and continued distress from the post-COVID transition to hybrid work environments.

The value destruction from the sudden shift to a high lease termination rate environment also erodes property tax revenues from commercial properties. Suppose the D.C. office market suffers the median expected valuation loss of $575 million. Applying the standard 1.89% statutory rate and the fact that assessment ratios are nearly 100% in D.C. results in local revenue losses of $51.26 million over a five-year period under a standard 3% risk-free discount rate applied to public cash flows (Congressional Budget Office, 2024).[38] Taken together, our results indicate that the economic consequences of underpriced risk from government lease contracts are not confined to directly affected landlords, but may propagate through capital markets, leasing dynamics, local services demand, and the property tax system. In this light, the DOGE lease terminations function not just as cost-cutting measures, but as systemic shocks with the potential to undermine CRE stability through both direct and indirect channels.

8 Conclusion

We offer new empirical evidence of how contractual risk embedded in federal lease agreements is priced in the commercial real estate debt market. We study the market's response to the federal government's large-scale exercise of early termination options (ETOs) during the 2025 policy shift led by the Department of Government Efficiency (DOGE). While ETOs had long existed as a legal clause in General Services Administration (GSA) leases, they were widely perceived as operationally dormant, rarely invoked, and thus ignored in risk assessments and bond pricing. The sudden wave of terminations in early 2025 represents a plausibly exogenous policy shock that reactivated this dormant contractual risk, providing a clear natural experiment to estimate the price of government contract risk premia.

[38] The Washington, D.C. Office of Tax and Revenue reports commercial property tax brackets and assessment ratios here: `https://otr.cfo.dc.gov/page/real-property-tax-rates`. Nearly all properties in our Trepp sample have an *ex ante* value greater than $10 million, corresponding to the 1.89% statutory tax bracket. This calculation assumes property value losses translate to lower tax assessments one-for-one in the first tax year and remain thereafter.

Leveraging proprietary data linking DOGE notifications to GSA lease records and Trepp CMBS values, we estimate difference-in-differences models using leases in first-loss tranches and not-yet ETO-eligible leases as a control group. We find that bond prices decline by 3.8%, and property-level NOI falls by 5.6% following ETO notification, relative to otherwise similar but non-notified leases. These effects are consistent with the theoretical mechanism highlighted by our arbitrage pricing framework for lease contingencies. We document large, negative spatial spillovers to the private-tenant segment of the market which are relatively constant in magnitude with respect to geographic proximity to terminated leases, pointing to a market-wide recalibration. Using our reduced form estimates of the decline in NOI for GSA and non-GSA leases stemming from early termination risk exposure, we simulate through the lens of our arbitrage pricing framework average expected property valuation losses of $575 million for the securitized mortgaged portion of the Washington, D.C. office market. Such losses dwarf the savings from canceled GSA lease payments for the D.C. metro area and likely negate total savings from nationwide lease terminations.

Our findings demonstrate that previously ignored contractual clauses – like the ETO – can become salient sources of credit and pricing risk once activated, and that federal policy shifts can generate significant valuation effects through this channel. Our results also call attention to the need for more explicit pricing of government contract risk in securitized credit products. In this setting, the DOGE intervention served as a wake-up call, revealing the latent exposure of CMBS structures to federal lease terminations.

As government leasing continues to evolve in the post-pandemic period, our analysis provides a foundational framework for understanding how public-sector behavior interacts with private capital markets through embedded contractual options. The negative risk exposure to government leases we document has the potential to magnify the troubles of regional banks suffering losses due to cratering commercial office valuations after the pandemic. As of 2021, 33% of agency and GSE-backed MBS investments were made by depository institutions (Fuster et al., 2025). Future work will explore implications of exposure to real estate policy uncertainty for the stability of the overall banking sector and credit provision.

References

Advan Research, 2022, Foot Traffic / Weekly Patterns [Dataset]. Dewey Data, https://doi.org/10.82551/X1PP-1F65.

Allen, Marcus T., Ronald C. Rutherford, and Larry J. Warner, 1997, A Comparison of Federal Government Office Rents with Market Rents, *Journal of Real Estate Finance and Economics* 15, 183–196.

Anenberg, Elliot, You Suk Kim, and Erica Moszkowski, 2025, Work-From-Home, Commercial Real Estate Risk and Credit Supply: Evidence from a Large Sample of Bank Loan Portfolios, *SSRN Working Paper* 5200882.

Ashcraft, Adam B., Kunal Gooriah, and Amir Kermani, 2019, Does Skin-In-The-Game Affect Security Performance, *Journal of Financial Economics* 134, 333–354.

Associated Press, 2025, Trump and Musk Demand Termination of Federal Office Leases Through General Services Administration.

Axios Local, 2024, Office buildings in D.C. Sit Empty As Sales Slump.

Baker, Scott R., Nicholas Bloom, and Steven J. Davis, 2016, Measuring Economic Policy Uncertainty, *The Quarterly Journal of Economics* 131, 1593–1636.

Baker, Scott R., and Constantine Yannelis, 2017, Income Changes and Consumption: Evidence from the 2013 Federal Government Shutdown, *Review of Economic Dynamics* 23, 99–124.

Benmelech, Efraim, Nittai Bergman, Anna Milanez, and Vladimir Mukharlyamov, 2018, The Agglomeration of Bankruptcy, *Review of Financial Studies* 32, 2541–2586.

Bertrand, Marianne, Esther Duflo, and Sendhil Mullainathan, 2004, How Much Should We Trust Differences-In-Differences Estimates?, *Quarterly Journal of Economics* 119, 249–275.

Bloomberg, 2025, DOGE's Lease Exits Put $12 Billion of Loans at Risk: Barclays.

BNP Paribas Real Estate, 2025, Market Research - Washington, D.C.

Brogaard, Jonathan, and Andrew Detzel, 2015, The Asset-Pricing Implications of Government Economic Policy Uncertainty, *Management Science* 61, 3–18.

Brown, Jordan, Ralf Meisenzahl, and Andy Polacek, 2024, Life Insurers' Exposure to Commercial Real Estate, *Federal Reserve Bank of Chicago Economic Perspectives* 5.

Callaway, Brantly, and Pedro H.C. Sant'Anna, 2021, Difference-in-Differences with Multiple Time Periods, *Journal of Econometrics* 225, 200–230.

CBS News, 2025, DOGE says it has canceled thousands of federal contracts, grants and leases. Here's a searchable list.

Chen, Yifan, Sean Wilkoff, and Jiro Yoshida, 2024, Amazon Is Coming to Town: Sequential Information Revelation in the Housing Market, *Real Estate Economics* 52, 277–323.

Cheng, Lin, Jacob Jaggi, Mark Yuzhi Yan, and Spencer Young, 2022, Debt Contracting and Changes to the Accounting for Leases: Implications of Accounting Standards Codification 842, *SSRN Working Paper* 4203733.

Choi, Soon Hyeok, Robert A. Jarrow, Daniel Lebret, and Crocker H. Liu, 2025, Pricing the Upside Potential to Downside Risk, *SSRN Working Paper* 5201008.

Christopoulos, Andreas D., Robert A. Jarrow, and Yildiray Yildirim, 2008, Commercial Mortgage-Backed Securities (CMBS) and Market Efficiency with Respect to Costly Information, *Real Estate Economics* 36, 441–498.

Cohn, Jonathan B., Zack Liu, and Malcolm I. Wardlaw, 2022, Count (and Count-Like) Data in Finance, *Journal of Financial Economics* 146, 529–551.

Col, Burcin, Art Durnev, and Alexander Molchanov, 2017, Foreign Risk, Domestic Problem: Capital Allocation and Firm Performance Under Political Instability, *Management Science* 64, 2102–2125.

Colliers Insights, 2018, GSA Lease Terminations: How Often Do They Occur?

CommercialEdge, 2025, CBD Office Markets Struggle Amid High Vacancies, Discounted Sales and Slowed Development.

Congressional Budget Office, 2024, How CBO Uses Discount Rates to Estimate the Present Value of Future Costs or Savings.

Conley, Timothy G., 2008, Spatial Econometrics, in Steven N. Durlauf, and Lawrence E. Blume, eds., *Microeconometrics*, 303–313 (London: Palgrave Macmillan).

CoStar News, 2025a, Federal Government Scraps Multiple Office Leases, Looks for More Across Country.

CoStar News, 2025b, US Government Has Now Canceled About One in 10 Federal Real Estate Leases.

Cresa, 2025, Washington, DC Market Report - Q2 2025.

Cushman & Wakefield, 2025, Washington, D.C. Office MarketBeat.

DC Office of Tax and Revenue, 2025, Tax Year 2025 Market Analytics Book.

de Chaisemartin, Clément, and Xavier D'Haultfoeuille, 2020, Two-Way Fixed Effects Estimators with Heterogeneous Treatment Effects, *American Economic Review* 110, 2964–2996.

Diamond, Rebecca, and Tim McQuade, 2019, Who Wants Affordable Housing in Their Backyard? An Equilibrium Analysis of Low-Income Property Development, *Journal of Political Economy* 127, 1063–1117.

Diener, Nicolas, Robert Jarrow, and Philip Protter, 2012, Relating Top-Down with Bottom-Up Approaches in the Evaluation of ABS with Large Collateral Pools, *International Journal of Theoretical and Applied Finance* 15, 1250011.

Duranton, Gilles, and William Kerr, 2018, Chapter 18: The Logic of Agglomeration, in *The New Oxford Handbook of Economic Geography*, 347–365 (Oxford University Press).

Federal Acquisition Regulation, 2025, Termination for Convenience of the Government (Fixed-Price).

Fitch Ratings, 2025, U.S. CMBS with Govt Lease Exposure Face Increased Default, Loss Risk.

Flynn, Sean, and Andra Ghent, 2018, Competition and Credit Ratings After the Fall, *Management Science* 64, 1672–1692.

Flynn, Sean, Andra C. Ghent, and Vasudha Nair, 2024, Determinants and Consequences of Return to Office Policies, *SSRN Working Paper* 4757876.

Fuster, Andreas, David Lucca, and James Vickery, 2025, Mortgage-Backed Securities, *FRB of Philadelphia Working Paper* 25-10.

Gelman, Michael, Shachar Kariv, Matthew D. Shapiro, Dan Silverman, and Steven Tadelis, 2020, How Individuals Respond to a Liquidity Shock: Evidence from the 2013 Government Shutdown, *Journal of Public Economics* 189, 103917.

Glancy, David, and J. Christina Wang, 2023, Lease Expirations and CRE Property Performance, *FRB of Boston Working Paper* 23-10.

Goldsmith-Pinkham, Paul, Peter Hull, and Michal Kolesár, 2024, Contamination Bias in Linear Regressions, *American Economic Review* 114, 4015–4051.

Goodell, John W., and Sami Vähämaa, 2013, US Presidential Elections and Implied Volatility: The Role of Political Uncertainty, *Journal of Banking & Finance* 37, 1108–1117.

Greenwald, Daniel, 2019, Firm Debt Covenants and the Macroeconomy: The Interest Coverage Channel, *SSRN Working Paper* 3535221.

Gupta, Arpit, Vrinda Mittal, and Stijn Van Nieuwerburgh, 2025, Work From Home and the Office Real Estate Apocalypse, *forthcoming, American Economic Review* .

Gupta, Arpit, Stijn Van Nieuwerburgh, and Constantine Kontokosta, 2022, Take the Q Train: Value Capture of Public Infrastructure Projects, *Journal of Urban Economics* 129.

Hassan, Tarek A., Stephan Hollander, Laurence van Lent, and Ahmed Tahoun, 2019, Firm-Level Political Risk: Measurement and Effects, *Quarterly Journal of Economics* 134, 2135–2202.

Herpfer, Christoph, Gonzalo Maturana, and Andrew Teodorescu, 2023, Intrinsic Motivation, Human Capital, and Productivity: Evidence from Government Shutdowns, *SSRN Working Paper* 3962279.

Holland & Knight, 2025, GSA Lease Terminations: Lessorsâ Rights and Responsibilities.

Hou, Young, Christopher Poliquin, Mariko Sakakibara, and Marco Testoni, 2025, Using Smartphone Location Data for Strategy Research, *SSRN Working Paper* 4796679.

Jarrow, Robert A., 2018, *Continuous-Time Asset Pricing Theory*, first edition (Springer International Publishing).

Jarrow, Robert A., 2021, The Economics of Insurance: A Derivatives-Based Approach, *Annual Review of Financial Economics* 13, 79–110.

Jiang, Eric Xuewei, Gregor Matvos, Tomasz Piskorski, and Amit Seru, 2024, Monetary Tightening and U.S. Bank Fragility in 2023: Mark-to-Market Losses and Uninsured Depositor Runs?, *Journal of Financial Economics* 159.

Jiang, Eric Xuewei, Gregor Matvos, Tomasz Piskorski, and Amit Seru, 2025, Monetary Tightening, Commercial Real Estate Distress, and US Bank Fragility, *forthcoming, Journal of Political Economy Macroeconomics* .

Julio, Brandon, and Youngsuk Yook, 2012, Political Uncertainty and Corporate Investment Cycles, *Journal of Finance* 67, 45–83.

Kelly, Bryan, Luboš Pástor, and Pietro Veronesi, 2016, The Price of Political Uncertainty: Theory and Evidence from the Option Market, *Journal of Finance* 71, 2417–2480.

Konishi, Hideo, and Michael T. Sandfort, 2003, Anchor Stores, *Journal of Urban Economics* 53, 413 – 435.

LaPoint, Cameron, 2022, Property Tax Sales, Private Capital, and Gentrification in the U.S., *SSRN Working Paper* 4219360.

Miyauchi, Yuhei, Kentaro Nakajima, and Stephen J. Redding, 2025, The Economics of Spatial Mobility: Theory and Evidence Using Smartphone Data, *forthcoming, Quarterly Journal of Economics* .

Mooradian, Robert M., and Shiawee X. Yang, 2000, Cancellation Strategies in Commercial Real Estate Leasing, *Real Estate Economics* 28, 65–88.

Newey, Whitney K., and Kenneth D. West, 1987, A Simple, Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix, *Econometrica* 55, 703–708.

Office of Management and Budget, 2024, OMB Report to Congress on Telework and Real Property Utilization.

Pashigian, B. Peter, and Eric D. Gould, 1998, Internalizing Externalities: The Pricing of Space in Shopping Malls, *Journal of Law and Economics* 41, 115–142.

Pasquariello, Paolo, and Christina Zafeiridou, 2014, Political Uncertainty and Financial Market Quality, *Michigan Ross School of Business Working Paper Series* 1232.

Pástor, Luboš, and Pietro Veronesi, 2012, Uncertainty about Government Policy and Stock Prices, *Journal of Finance* 67, 1219–1264.

Pástor, Luboš, and Pietro Veronesi, 2013, Political Uncertainty and Risk Premia, *Journal of Financial Economics* 110, 520–545.

Politico, 2025, Inside GSAâs Mass Lease Terminations.

Qian, Franklin, Qianyang Zhang, and Xiang Zhang, 2024, Identifying Agglomeration Spillovers: Evidence from Grocery Store Openings, Technical report, UNC Chapel Hill.

Rambachan, Ashesh, and Jonathan Roth, 2023, A More Credible Approach to Parallel Trends, *Review of Economic Studies* 90, 2555–2591.

Rosenthal, Stuart S., and William C. Strange, 2020, How Close Is Close? The Spatial Reach of Agglomeration Economies, *Journal of Economic Perspectives* 34, 27–49.

Roth, Jonathan, 2022, Pretest with Caution: Event-Study Estimates after Testing for Parallel Trends, *American Economic Review: Insights* 4, 305–322.

Saiegh, Sebastian M., 2023, Fear and Loathing on Wall Street: Electoral Risk Hedging in the United States (1986-2020), *Journal of Political Economy and Political Institutions* 4, 311–331.

Sakong, Jung, and Alexander Zentefis, 2025, What Drives Disparities in Bank Branch Use? A Geospatial Analysis, *SSRN Working Paper* 4621778.

Trepp Research, 2025, DOGE Cuts Begin on GSA-Leased Office Space: February Data Update.

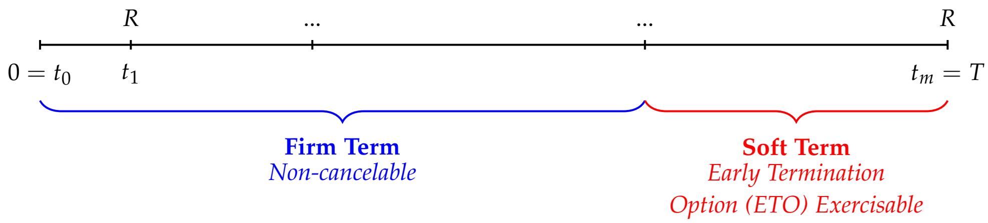
U.S. General Services Administration, 2023, Lease Offer Submission Guide (LOSG).

U.S. General Services Administration, 2024, GSA Lease Policies and Templates.

Yahoo Finance, 2025, Government Looks To Cut 100 Federal Leases In D.C. Causing Market Disruption.

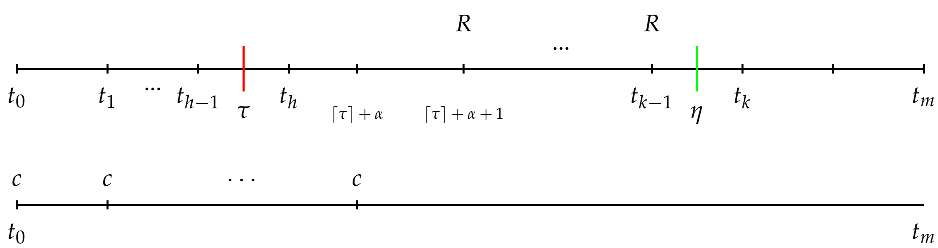
Figures

FIGURE 1. Federal Lease Term Structure with an Early Termination Option (ETO)



Notes: The figure plots the timing of a typical lease term divided into a firm term and a soft term. During the firm term, the lease is non-cancelable and the early termination option is unavailable. In the soft term, the tenant may exercise the early termination option. The timeline lengths reflect that for the median federal lease contract, the firm term lasts ten years, and the soft term lasts five years.

FIGURE 2. Timeline of Key Events with an Exercised ETO



Notes: The figure plots the sequence of key events following an exercised ETO during the soft term of a federal lease. In the above timeline, the initial event, denoted by τ, marks the time at which the federal tenant issues formal notice of its intent to terminate, subject to the required advance notice period α. Upon expiration of this notice period, the landlord assumes responsibility for the vacated rent obligation until a replacement tenant is secured at a subsequent date, denoted by η. The bottom timeline illustrates that federal tenants pay an insurance premium c for holding a long position on their ETO.

FIGURE 3. Termination Right and Rent Abatement Clauses for GSA Leases

A. Termination Rights and (Non-)Renewal Clauses

1.05 TERMINATION RIGHTS (OCT 2016)

The Government may terminate this Lease, in whole or in parts, at any time effective after the Firm Term of this Lease, by providing not less than **XX** days' prior written notice to the Lessor. The effective date of the termination shall be the day following the expiration of the required notice period or the termination date set forth in the notice, whichever is later. No rental shall accrue after the effective date of termination.

1.06 RENEWAL RIGHTS (OCT 2016)

A. This Lease may be renewed at the option of the Government for a term of **XX YEARS** at the following rental rate(s):

	OPTION TERM, YEARS XX - XX	
	ANNUAL RENT	ANNUAL RATE / RSF
SHELL RENTAL RATE	$XX	$XX
OPERATING COSTS	OPERATING COST BASE SHALL CONTINUE FROM THE EFFECTIVE YEAR OF THE LEASE. OPTION TERM IS SUBJECT TO CONTINUING ANNUAL ADJUSTMENTS.	

provided notice is given to the Lessor at least **XX** days before the end of the original Lease term or any extension thereof; all other terms and conditions of this Lease, as same may have been amended, shall remain in full force and effect during any renewal term.

NOTE: REVISE SUB-PARAGRAPH B IF THE INTENT IS TO SEEK FIRM TERM RENEWAL OPTIONS.

B. Termination rights outlined in the "Termination Rights" paragraph apply to all renewal terms.

B. Rent Abatement Clauses

ACTION REQUIRED: USE IF THERE IS A NEGOTIATED AMOUNT FOR THE VACANT LEASED PREMISES.

NOTE: ALWAYS ATTEMPT TO NEGOTIATE AN ADJUSTMENT FOR VACANT PREMISES PRIOR TO LEASE AWARD. IDEALLY, NEGOTIATE OUT ALL NON-REQUIRED SERVICES AND UTILITIES IN THE VACANT SPACE.

1.15 RATE FOR ADJUSTMENT FOR VACANT LEASED PREMISES (SEP 2013)

In accordance with the paragraph entitled "Adjustment for Vacant Premises," if the Government fails to occupy or vacates the entire or any portion of the Premises prior to expiration of the term of the Lease, the operating costs paid by the Government as part of the rent shall be reduced by **$XX.XX** per ABOA SF of Space vacated by the Government.

NOTE: ALWAYS ATTEMPT TO NEGOTIATE SOME KIND OF ADJUSTMENT FOR VACANT PREMISES PRIOR TO LEASE AWARD. IDEALLY, NEGOTIATE OUT ALL NON-REQUIRED SERVICES AND UTILITIES IN THE VACANT SPACE.

2.08 GSAR 552.270-16 ADJUSTMENT FOR VACANT PREMISES (DEVIATION) (SEP 2022)

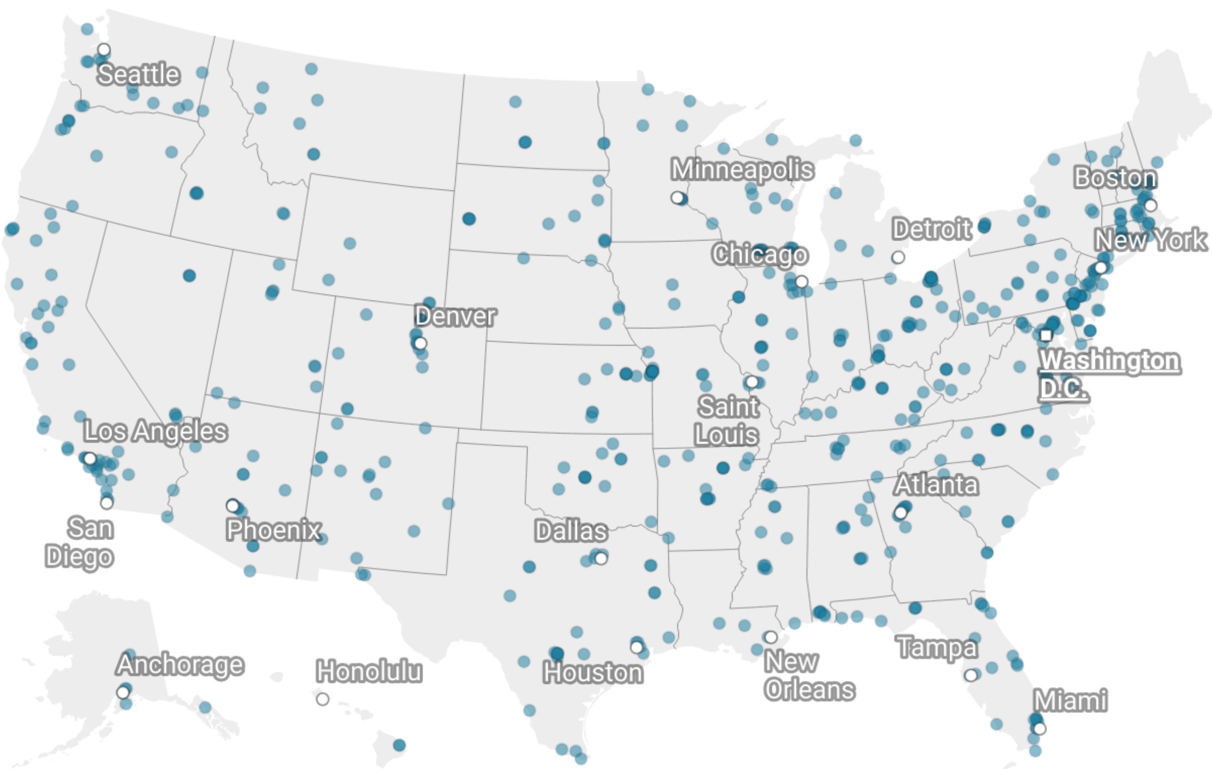
(a) If the Government fails to occupy any portion of the leased premises or vacates the premises in whole or in part prior to expiration of the term of the lease, the rental rate and the base for operating cost adjustments will be reduced using the figure specified in the "Rate for Adjustment for Vacant Leased Premises" paragraph of this Lease.

(b) If no rate reduction has been established in this lease, the rate will be reduced by that portion of the costs per ABOA square foot of operating expenses not required to maintain the space.

(c) Said reduction shall occur after the Government gives 30 calendar days' prior notice to the Lessor and shall continue in effect until the Government occupies the vacant premises or the lease expires or is terminated.

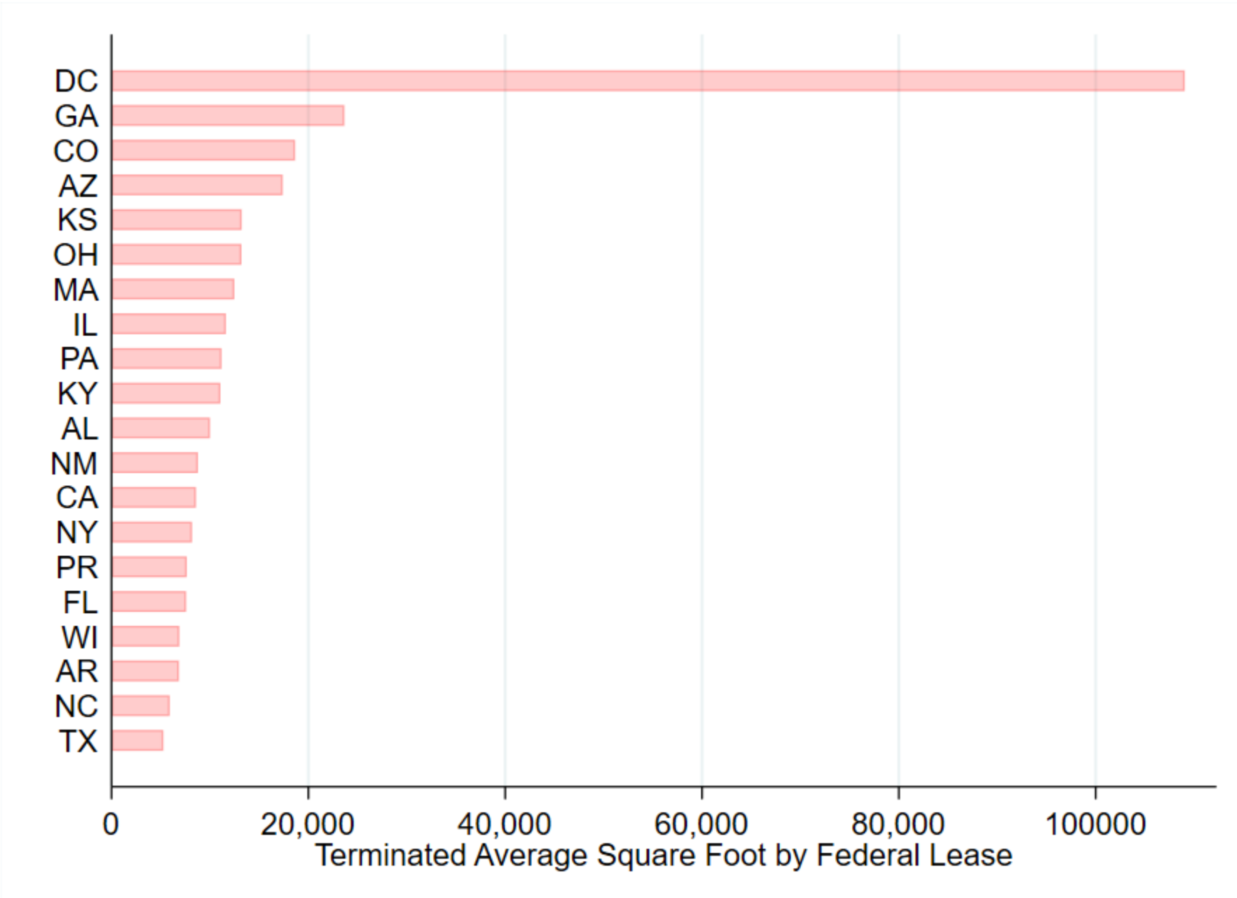
Notes: The figure consists of sample GSA termination right clauses in Panel A (§1.05 and §1.06) and the vacant premises clauses in Panel B (§ 1.15 and § 2.08). The §1.05 clause stipulates the terms of the government's early termination option (ETO). §1.06 indicates that, by default, termination rights apply to any renewal terms as well. The §1.15 clause provides a rent abatement if the federal tenant vacates the premise partially at any time before its expiration, including in the event that the government exits during the soft term of the lease by exercising its ETO. §2.08 clarifies that if the GSA and private contractor did not agree to rent abatement rate in §1.15, then the rent abatement is determined by the fraction of operating expenditures attributable to the leasable square footage (ABOA) occupied by the GSA tenant. ANSI/BOMA Office Area (ABOA) means the area "where a tenant normally houses personnel, and/or furniture, for which a measurement is to be computed," as stated by the American National Standards Institute/Building Owners and Managers Association (ANSI/BOMA) publication, Z65.1-1996. *Source*: GSA Global Lease Template L100, revised October 2023.

FIGURE 4. Federal Lease Termination Footprint by State



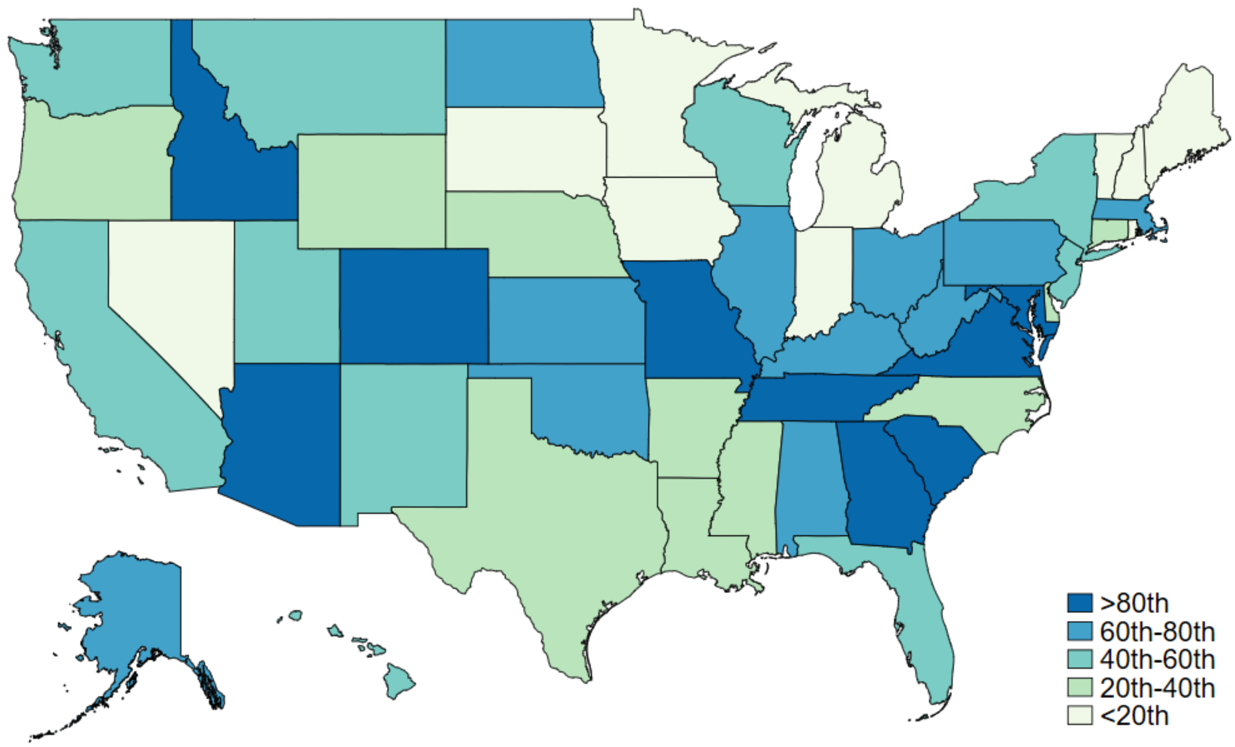
Notes: The map plots a snapshot of the DOGE-terminated federal leases as of March 24, 2025 using Datawrapper. On March 24, 2025, the number of terminated leases reached 679. *Sources*: Department of Government Efficiency, Arco Real Estate Solutions, JLL Federal Lease Termination Tracker.

FIGURE 5. Top 20 States in Terminated Average Square Footage by Federal Lease



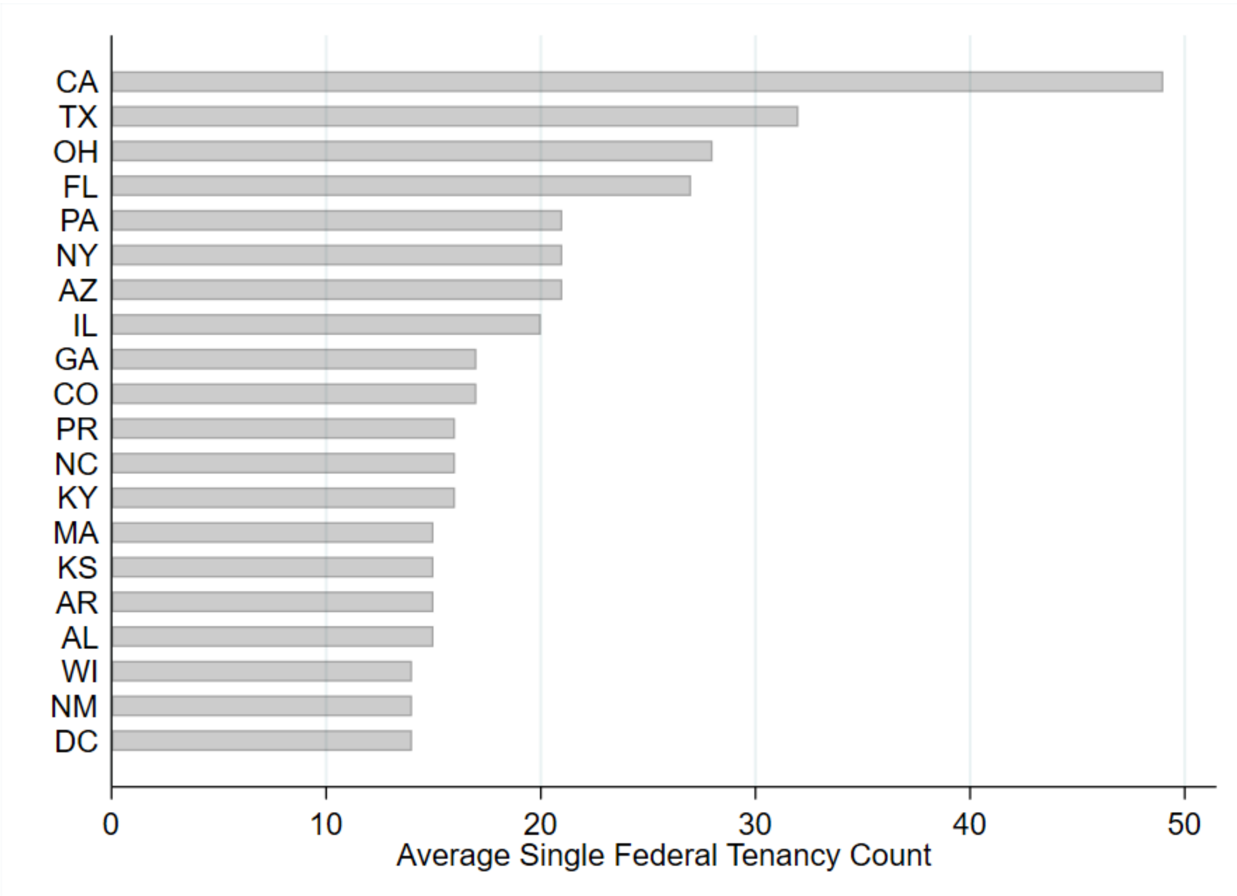
Notes: The bar plot shows the average square foot of terminated federal leases of top 20 states in descending order as of March 24, 2025. *Source*: Department of Government Efficiency, Arco Real Estate Solutions, JLL Federal Lease Termination Tracker.

FIGURE 6. Fraction of Total Square Footage Terminated by State



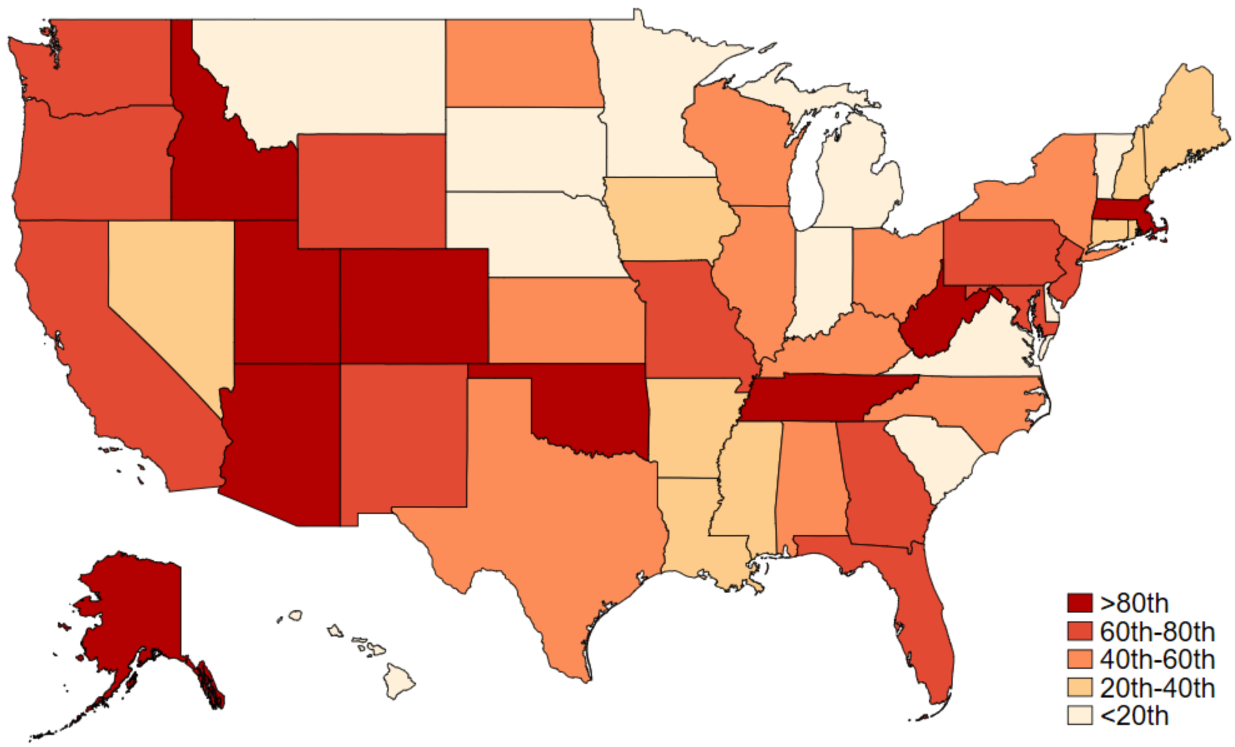
Notes: The map displays the fraction of total square footage in terminated federal leases by state in quintiles as of March 24, 2025. *Sources*: Department of Government Efficiency, Arco Real Estate Solutions, JLL Federal Lease Termination Tracker.

FIGURE 7. Top 20 States in Single-Tenancy Federal Lease Concentration



Notes: The bar plot shows the average number of single-tenancy federal leases of top 20 states in descending order as of March 24, 2025. *Sources*: Department of Government Efficiency, Arco Real Estate Solutions, JLL Federal Lease Termination Tracker.

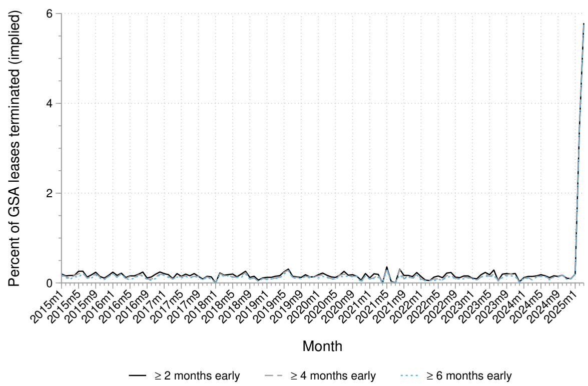
FIGURE 8. Fraction of Total Savings due to Terminated Federal Leases by State



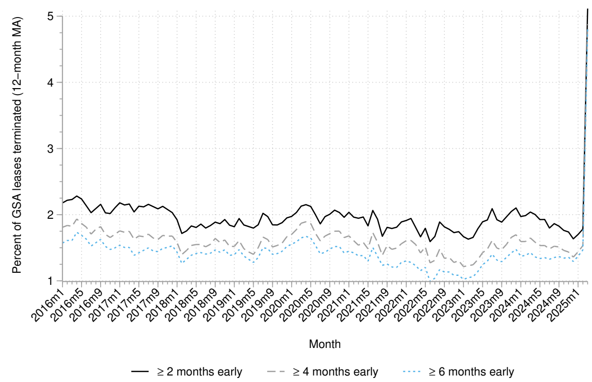
Notes: The map displays the fraction of total savings reported by the Department of Government Efficiency (DOGE) due to federal lease terminations by state in quintiles as of March 24, 2025. *Source*: Department of Government Efficiency, Arco Real Estate Solutions, JLL Federal Lease Termination Tracker.

FIGURE 9. GSA Lease Cancellation Rates Implied by GSA Inventory and DOGE Announcements

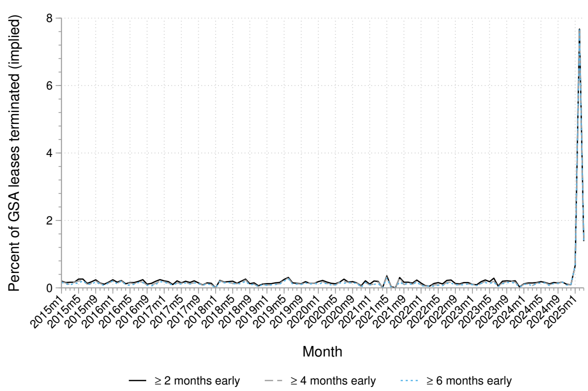
A. Monthly Rates: DOGE Website



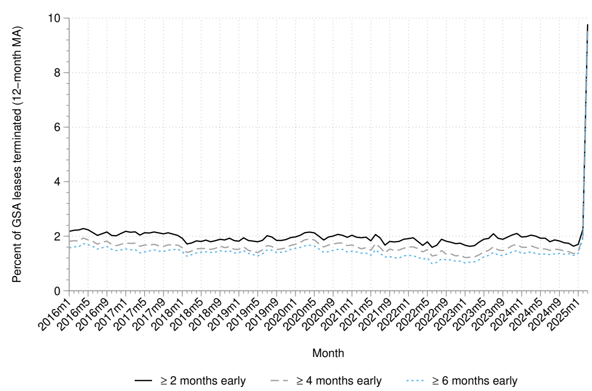
B. 12-Month Rates: DOGE Website



C. Monthly Rates: DOGE Letter



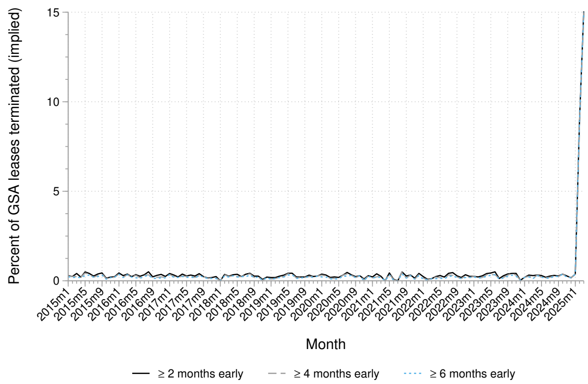
D. 12-Month Rates: DOGE Letter



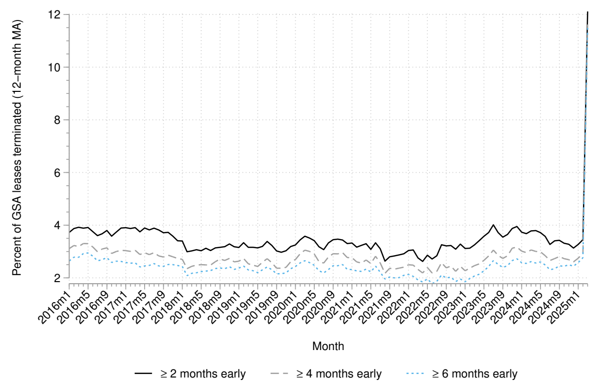
Notes: The figure plots measures of the empirical probability of a GSA lease being canceled before it expires. We plot the monthly termination rate in the left-hand side panels (Panels A and C), whereas Panels B and D plot the 12-month lagged sum of monthly termination rates to account for seasonality. We follow the procedures outlined by Colliers Insights (2018) to identify canceled leases from the historical GSA inventory lists. In particular, we drop leases which were superseded in the same property, or those that were short term or temporary leases which were originally executed with less than three years of firm term (i.e., pre-ETO eligibility period) or those with less than five years in the total term. We keep all leases in the sample regardless of their rentable square footage. We flag a lease as canceled and mark the last month it appears in the GSA inventory panel if it disappears from the panel at least x months prior to lease expiration, where for robustness we vary $x = \{2, 4, 6\}$ months in the figures above. For the months starting in January 2025 when DOGE was created, we add in the implied terminations due to either the timing of leases listed on the DOGE website (Panels A and B) or the timing of the notifications sent by DOGE to the tenants and landlords (Panels C and D). The series are therefore "implied" cancellation rates, because some leases were removed from the DOGE termination list after March 2025. *Sources*: Department of Government Efficiency, Inventory of GSA Owned and Leased Properties.

FIGURE 10. ETO-Eligible Cancellation Rates Implied by GSA Inventory and DOGE Announcements

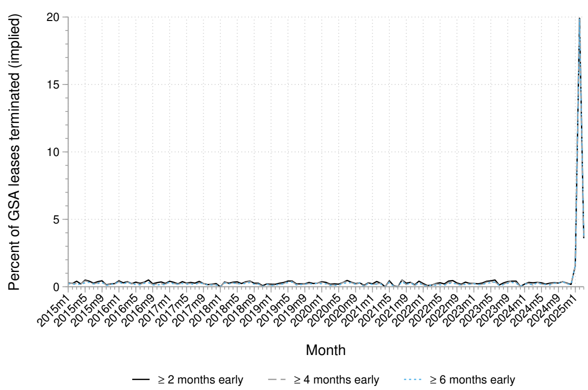
A. Monthly Rates: DOGE Website



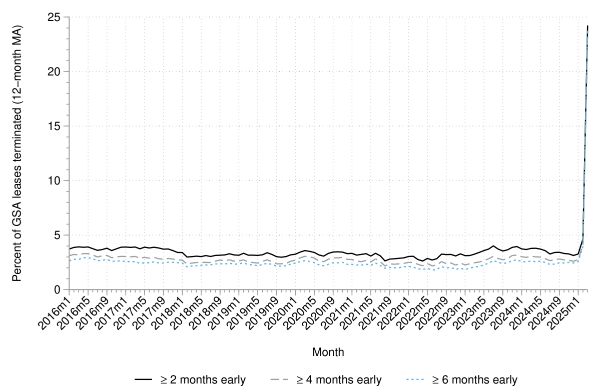
B. 12-Month Rates: DOGE Website



C. Monthly Rates: DOGE Letter

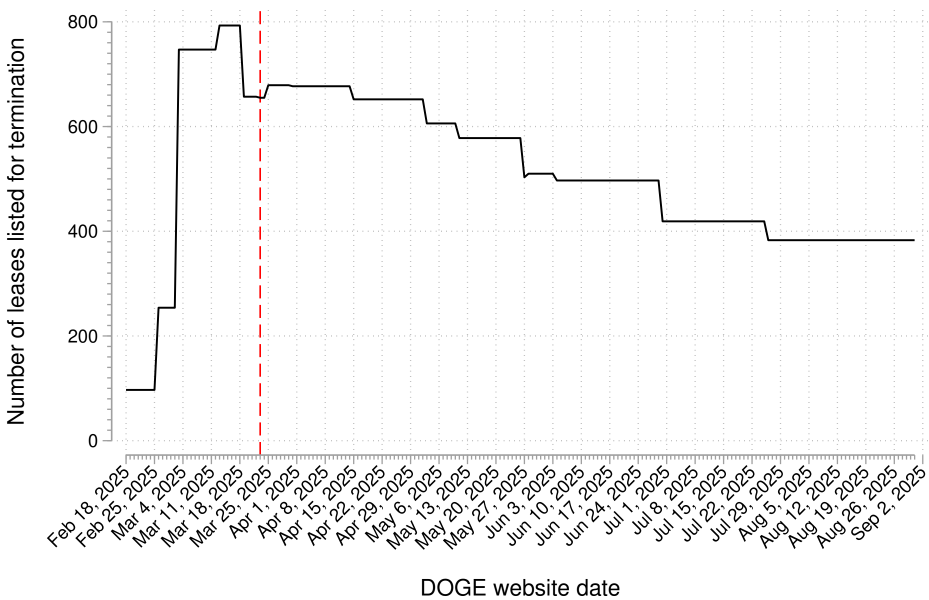


D. 12-Month Rates: DOGE Letter



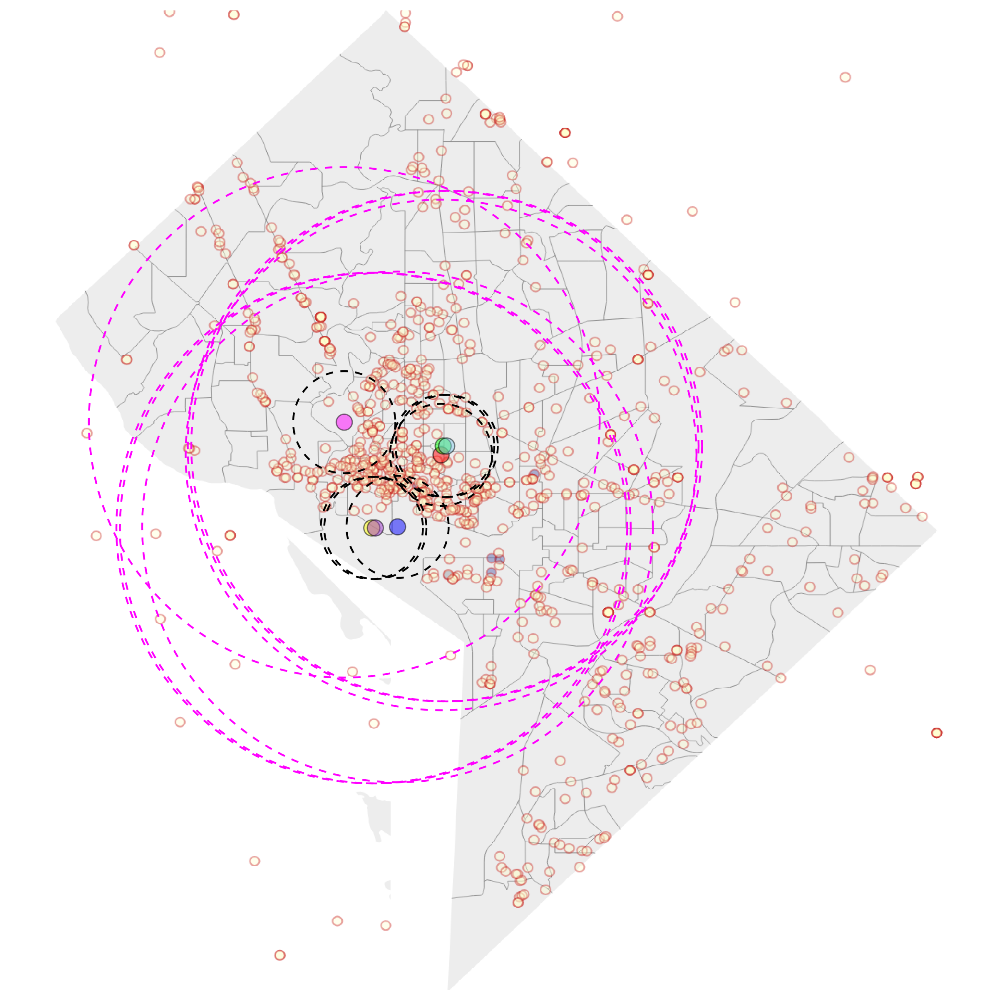
Notes: The figure plots measures of the empirical probability of a GSA lease being canceled before it expires. Relative to Figure 9, the procedures are identical except we now restrict attention to ETO-eligible leases, meaning those which are in the soft term where the date is after the termination right date. We plot the monthly termination rate in the left-hand side panels (Panels A and C), whereas Panels B and D plot the 12-month lagged sum of monthly termination rates to account for seasonality. We follow the procedures outlined by Colliers Insights (2018) to identify canceled leases from the historical GSA inventory lists. In particular, we drop leases which were superseded in the same property, or those that were short term or temporary leases which were originally executed with less than three years of firm term (i.e., pre-ETO eligibility period) or those with less than five years in the total term. We keep all leases in the sample regardless of their rentable square footage. We flag a lease as canceled and mark the last month it appears in the GSA inventory panel if it disappears from the panel at least x months prior to lease expiration, where for robustness we vary $x = \{2, 4, 6\}$ months in the figures above. For the months starting in January 2025 when DOGE was created, we add in the implied terminations due to either the timing of leases listed on the DOGE website (Panels A and B) or the timing of the notifications sent by DOGE to the tenants and landlords (Panels C and D). The series are therefore "implied" cancellation rates, because some leases were removed from the DOGE termination list after March 2025. *Sources*: Department of Government Efficiency, Inventory of GSA Owned and Leased Properties.

FIGURE 11. Number of Canceled GSA Leases Announced by DOGE: February 18, 2025 – August 31, 2025



Notes: We plot the number of GSA leases listed as terminated on the DOGE lease savings website since the first date the information was published on February 18, 2025 up until August 31, 2025. The number of canceled leases peaked at 793 between March 13, 2025 and March 18, 2025. The dashed vertical red line indicates the date (March 23, 2025) that we started scraping the DOGE website. Information on canceled leases prior to March 23, 2025 obtained from the Wayback Machine and various real estate news sources. *Sources*: Department of Government Efficiency, Arco Real Estate Solutions, JLL Federal Lease Termination Tracker.

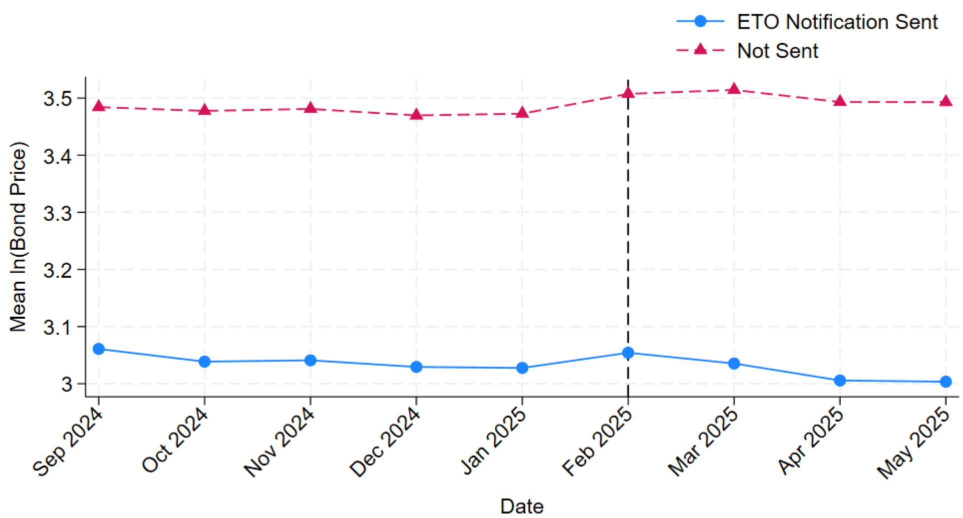
FIGURE 12. A 1-Mile Inner Ring and 5-Mile Outer Ring around Terminated Washington, D.C. Lease Properties



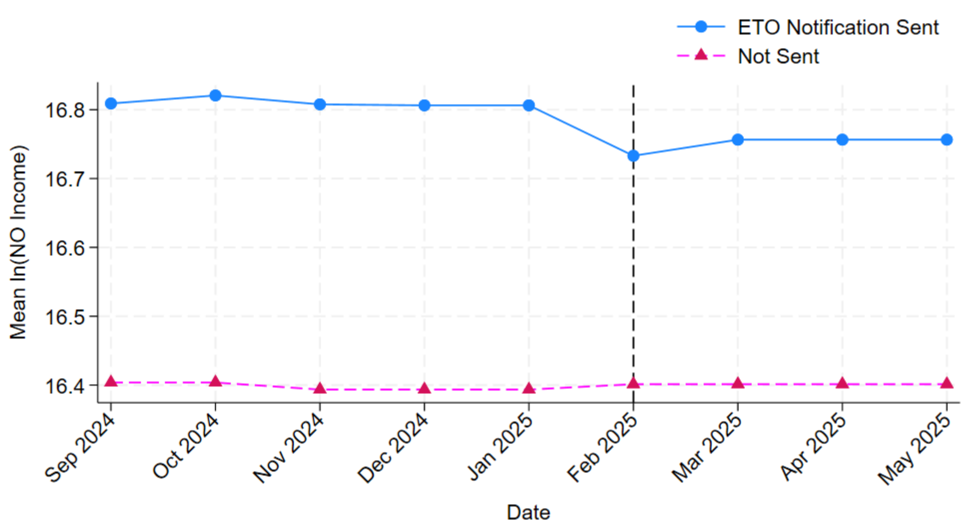
Notes: The map plots a set of 1-mile rings (black) around the seven terminated Washington, D.C. federal lease properties tied to a CMBS deal (with a 5-mile buffer, colored in purple) using Datawrapper. The ring radii correspond to the baseline parametrizations of our spatial difference-in-differences specification, as outlined in Section 5.3. Each colored, filled colored circles inside the black rings on the map represents a DOGE-terminated GSA lease which received formal notification. Yellow points indicate other securitized properties in the Trepp data which were not notified by DOGE. We include properties in the Washington, D.C. MSA, which includes counties in Maryland and Virginia. *Sources*: Department of Government Efficiency.

FIGURE 13. First-Loss Group (FLG) Tranches: Average Bond Prices and NOI Time Series by ETO Notification Status

A. Log CMBS Bond Prices



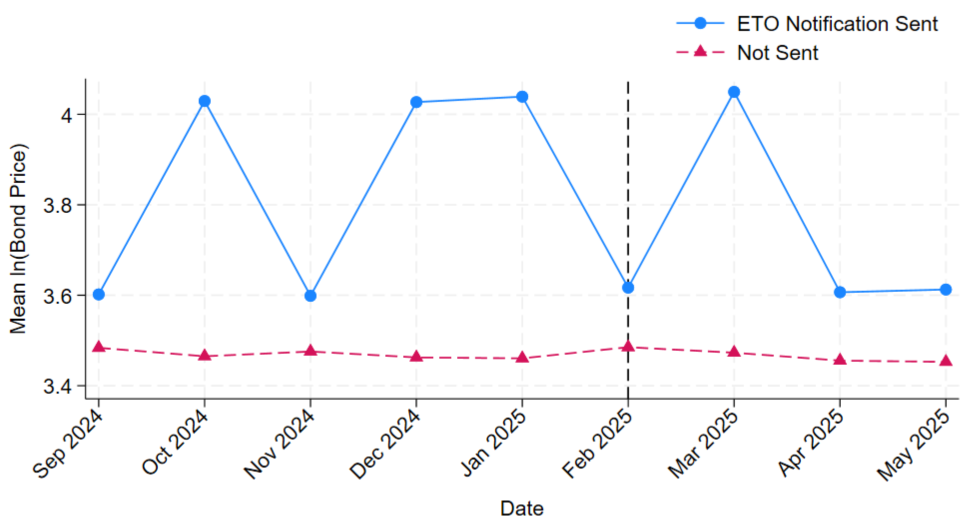
B. Log Net Operating Income (NOI) for Properties with ETO-Eligible Leases



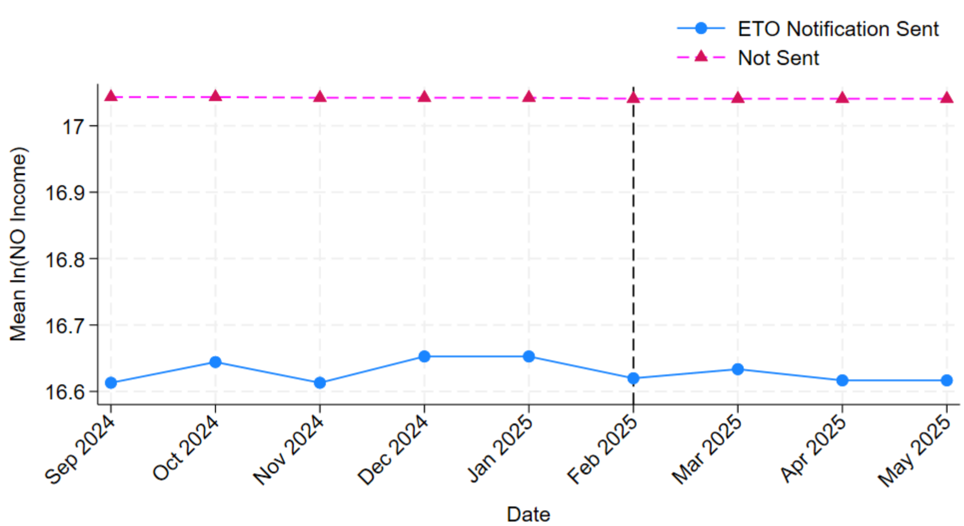
Notes: Panel A of the figure plots the time series of average bond prices for first-loss group (FLG) tranches over time which are linked to Washington, D.C. leases eligible for the early termination option (ETO). The blue series represents bond prices attached to leases with ETO notification sent by the government, while the pink series represents bonds attached to our control group of leases which did not receive a cancellation notification because they were not-yet eligible. Panel B repeats the exercise for log property NOI for properties associated with ETO-eligible leases included in FLG tranches. We follow Flynn and Ghent (2018) in defining the FLG as consisting of tranches which have a rating of CCC, or CCC+, or are unrated. In cases where the bond receives multiple agency ratings, we use the S&P rating. If the S&P rating is unavailable, we use the Fitch rating. Finally, if both the S&P and Fitch ratings are unavailable, we adopt the Moody's rating. We winsorize bond prices at the 1st and 99th percentiles and winsorize NOI at the median $\pm$ 5 times the interquartile range.

FIGURE 14. Mezzanine Tranches: Average Bond Prices and NOI Time Series by ETO Notification Status

A. Log CMBS Bond Prices



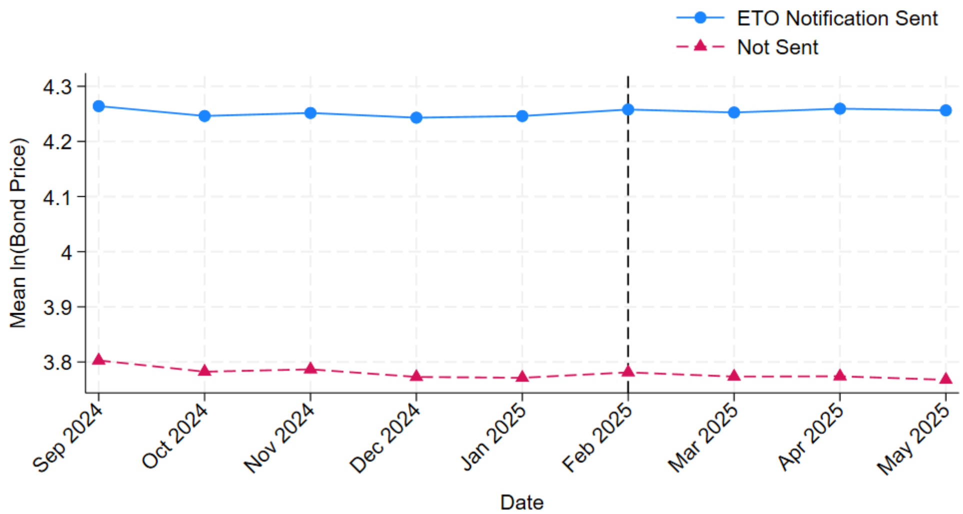
B. Log Net Operating Income (NOI) for Properties with ETO-Eligible Leases



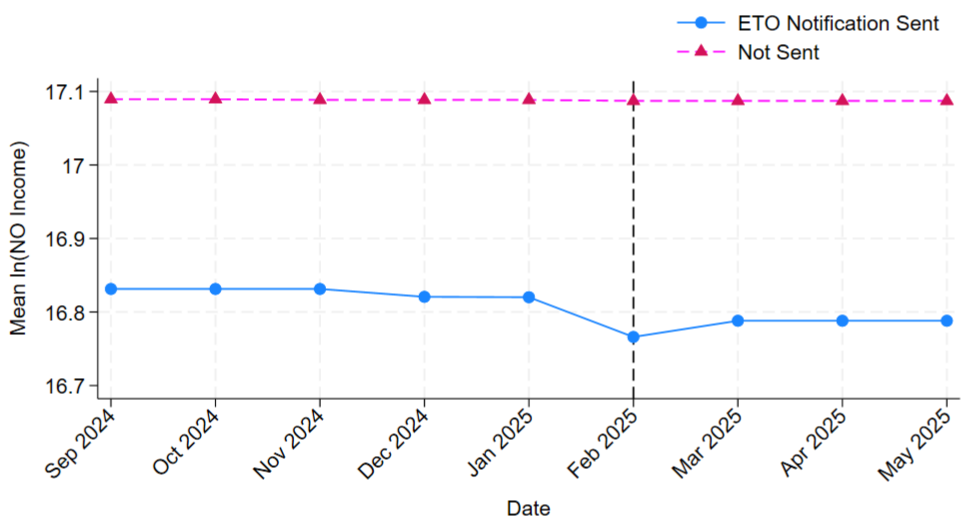
Notes: The figure plots the time series of average bond prices for mezzanine tranches over time which are linked to Washington, D.C. leases eligible for the early termination option (ETO). The blue series represents bond prices attached to leases with ETO notification sent by the government, while the pink series represents bonds attached to our control group of leases which did not receive a cancellation notification because they were not-yet eligible. Panel B repeats the exercise for log property NOI for properties associated with ETO-eligible leases included in mezzanine tranches. We follow Flynn and Ghent (2018) in defining the mezzanine set of tranches as those rated below AAA but above CCC+. In cases where the bond receives multiple agency ratings, we use the S&P rating. If the S&P rating is unavailable, we use the Fitch rating. Finally, if both the S&P and Fitch ratings are unavailable, we adopt the Moody's rating. We winsorize bond prices at the 1st and 99th percentiles and winsorize NOI at the median $\pm$ 5 times the interquartile range.

FIGURE 15. Senior Tranches: Average Bond Prices and NOI Time Series by ETO Notification Status

A. Log CMBS Bond Prices

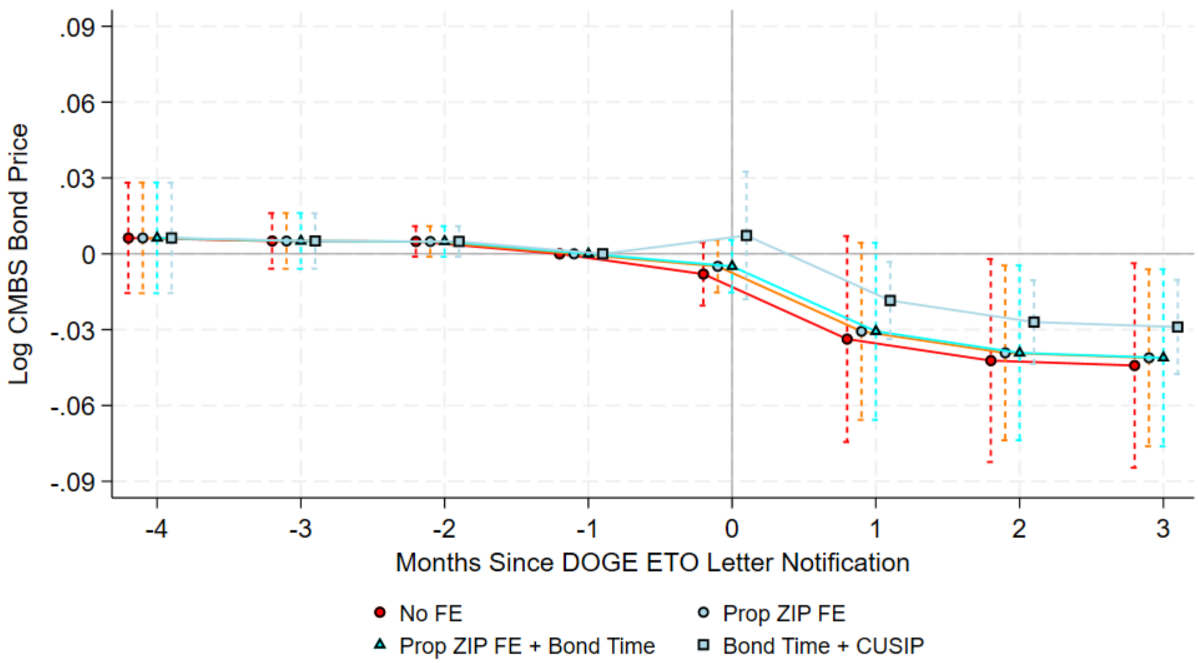


B. Log Net Operating Income (NOI) for Properties with ETO-Eligible Leases



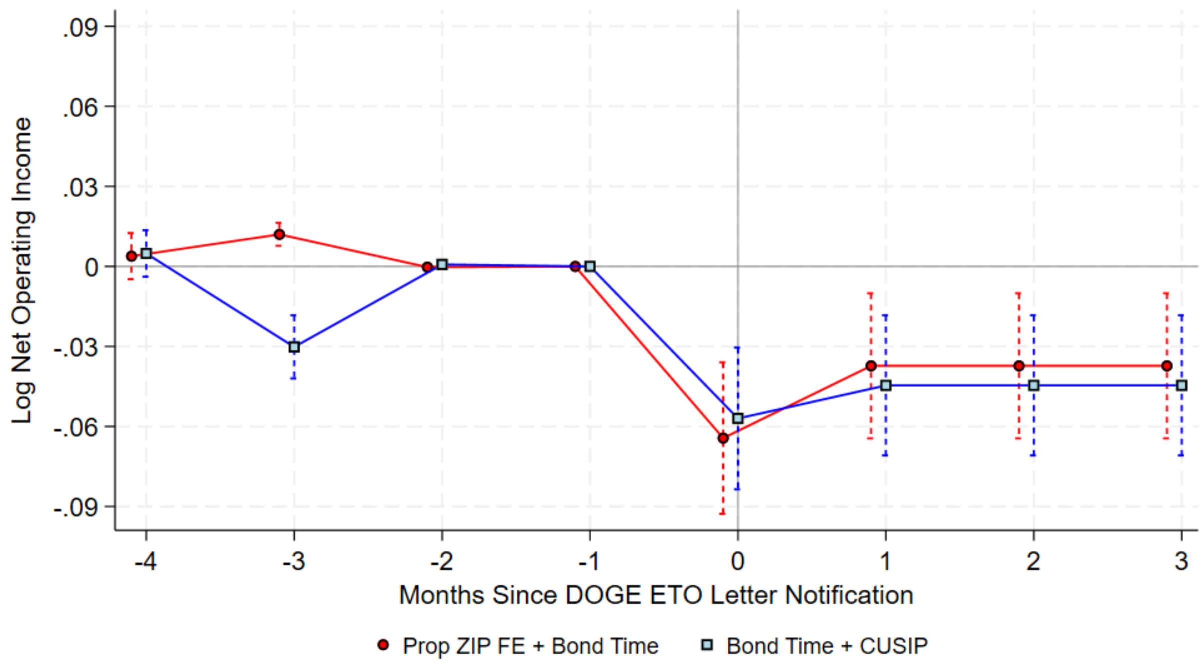
Notes: The figure plots the time series of average bond prices for senior tranches over time which are linked to Washington, D.C. leases eligible for the early termination option (ETO). The blue series represents bond prices attached to leases with ETO notification sent by the government, while the pink series represents bonds attached to our control group of leases which did not receive a cancellation notification because they were not-yet eligible. Panel B repeats the exercise for log property NOI for properties associated with ETO-eligible leases included in senior tranches. We follow Flynn and Ghent (2018) in defining the senior set of tranches as those rated AAA. In cases where the bond receives multiple agency ratings, we use the S&P rating. If the S&P rating is unavailable, we use the Fitch rating. Finally, if both the S&P and Fitch ratings are unavailable, we adopt the Moody's rating. We winsorize bond prices at the 1st and 99th percentiles and winsorize NOI at the median $\pm$ 5 times the interquartile range.

FIGURE 16. Dynamic Diff-in-Diff Plot: SE Cluster at Bond CUSIP Level
(Y: Log CMBS Bond Price, Tranche: First-Loss Group)



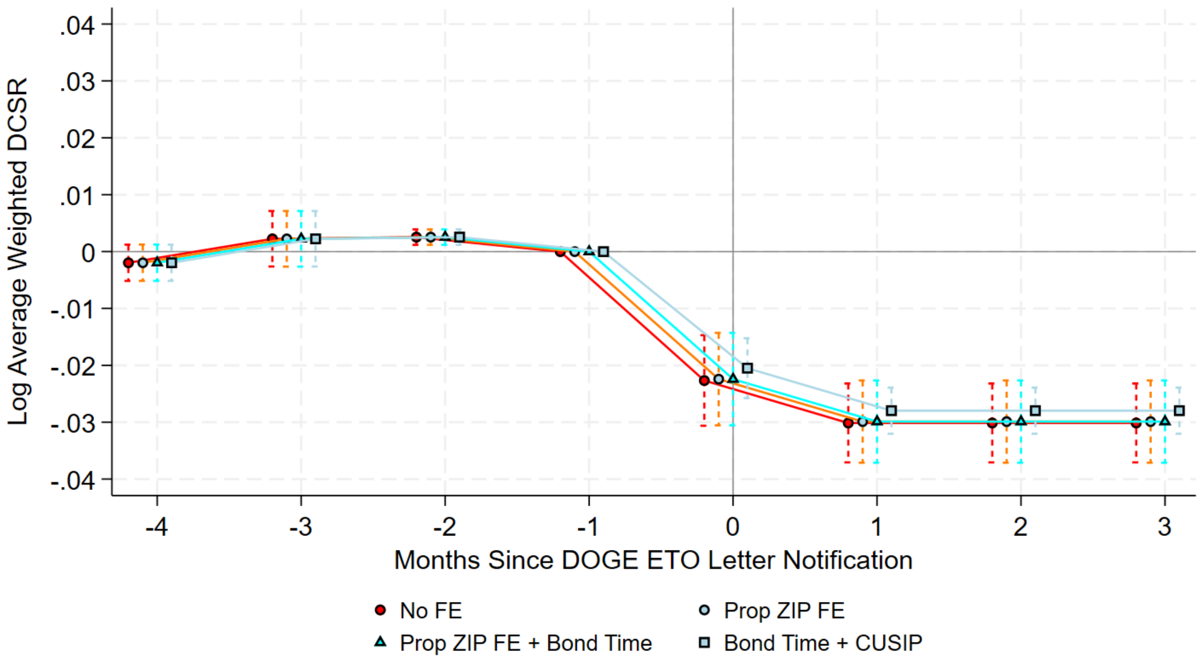
Notes: The figure plots the event study coefficients estimated from versions of equation (5.2) with log CMBS bond price as the outcome variable, applied to a 4-month symmetric window around the creation of DOGE in January 2025. We plot the estimates from four specifications: one without any fixed effects ("No FE"), then one adding in property 5-digit zip code fixed effects ("Prop ZIP FE"), then adding deal date fixed effects ("Prop ZIP FE + Bond Time"), then finally adding CUSIP fixed effects ("Bond Time + CUSIP FE"). In each specification, we set $t = -1$ to be the reference period, corresponding to January 2025. This reference period choice reflects the fact that the first set of DOGE terminations for the Washington, D.C. area was sent to landlords and tenants on January 30, 2025. We restrict our sample to the first-loss group of tranches. We continue to follow Flynn and Ghent (2018) in defining the FLG as consisting of tranches which have a rating of CCC, or CCC+, or are unrated. We winsorize bond prices at the 1st and 99th percentiles. 95% confidence intervals obtained from clustering standard errors at the bond CUSIP level.

FIGURE 17. Dynamic Diff-in-Diff Plot: SE Cluster at Bond CUSIP Level
(Y: Log NOI, Tranche: First-Loss Group)



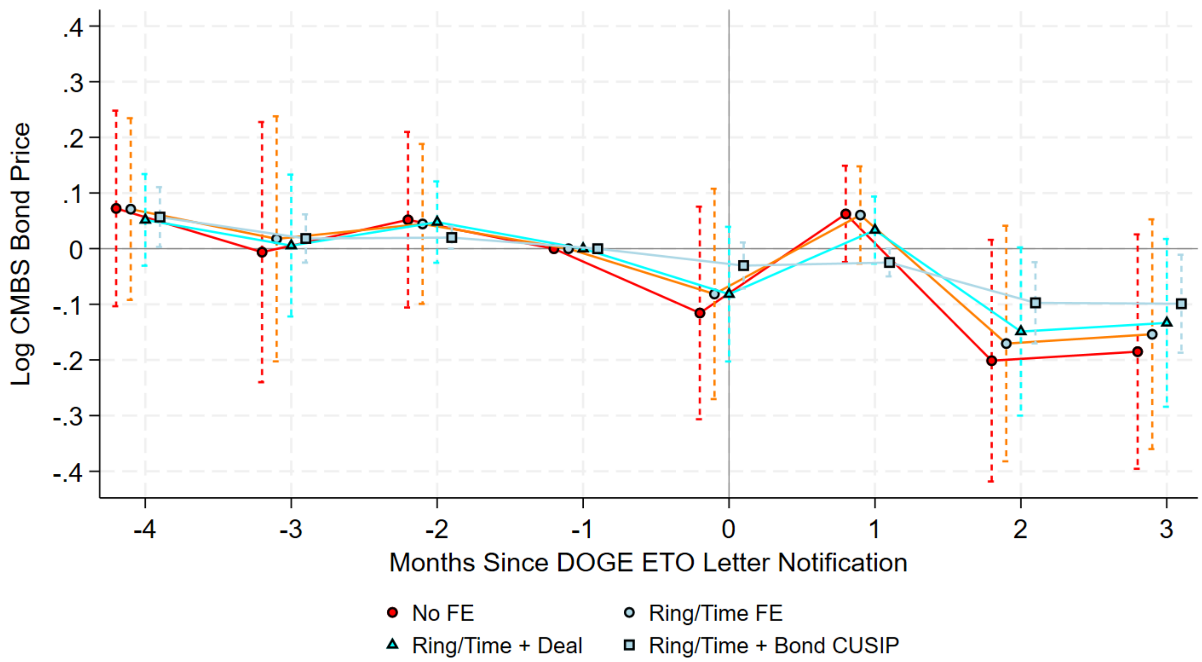
Notes: The figure plots the event study coefficients estimated from versions of equation (5.2) with log property net operating income as the outcome variable, applied to a 4-month symmetric window around the creation of DOGE in January 2025. We plot the estimates from two specifications: one with property 5-digit zip codes and deal date fixed effects ("Prop ZIP FE + Bond Time") and another with deal date and CUSIP fixed effects ("Bond Time + CUSIP FE"). We measure NOI in the Trepp data as the most recent NOI attached to a securitized loan. In each specification, we set $t = -1$ to be the reference period, corresponding to January 2025. This reference period choice reflects the fact that the first set of DOGE terminations for the Washington, D.C. area was sent to landlords and tenants on January 30, 2025. We restrict our sample to properties with mortgages pooled in the first-loss group of tranches. We continue to follow Flynn and Ghent (2018) in defining the FLG as consisting of tranches which have a rating of CCC, or CCC+, or are unrated. We winsorize NOI at the median ± 5 times the interquartile range. 95% confidence intervals obtained from clustering standard errors at the bond CUSIP level.

FIGURE 18. Dynamic Diff-in-Diff Plot: SE Cluster at Bond CUSIP Level
(Y: Log Avg Weighted DSCR, Tranche: First-Loss Group)



Notes: The figure plots the event study coefficients estimated from versions of equation (5.2) with log weighted average DSCR as the outcome variable, applied to a 4-month symmetric window around the creation of DOGE in January 2025. We plot the estimates from four specifications: one without any fixed effects ("No FE"), then one adding in property 5-digit zip code fixed effects ("Prop ZIP FE"), then adding deal date fixed effects ("Prop ZIP FE + Bond Time"), then finally adding CUSIP fixed effects ("Bond Time + CUSIP FE"). In each specification, we set $t = -1$ to be the reference period, corresponding to January 2025. This reference period choice reflects the fact that the first set of DOGE terminations for the Washington, D.C. area was sent to landlords and tenants on January 30, 2025. We restrict our sample to the first-loss group of tranches. The weighted DSCR is defined as the share-weighted average across properties in the loan pool of the ratio of property NOI to debt service. We measure DSCR in Trepp as of the bond deal distribution date. We continue to follow Flynn and Ghent (2018) in defining the FLG as consisting of tranches which have a rating of CCC, or CCC+, or are unrated. We winsorize NOI at the median $\pm$ 5 times the interquartile range. 95% confidence intervals obtained from clustering standard errors at the bond CUSIP level.

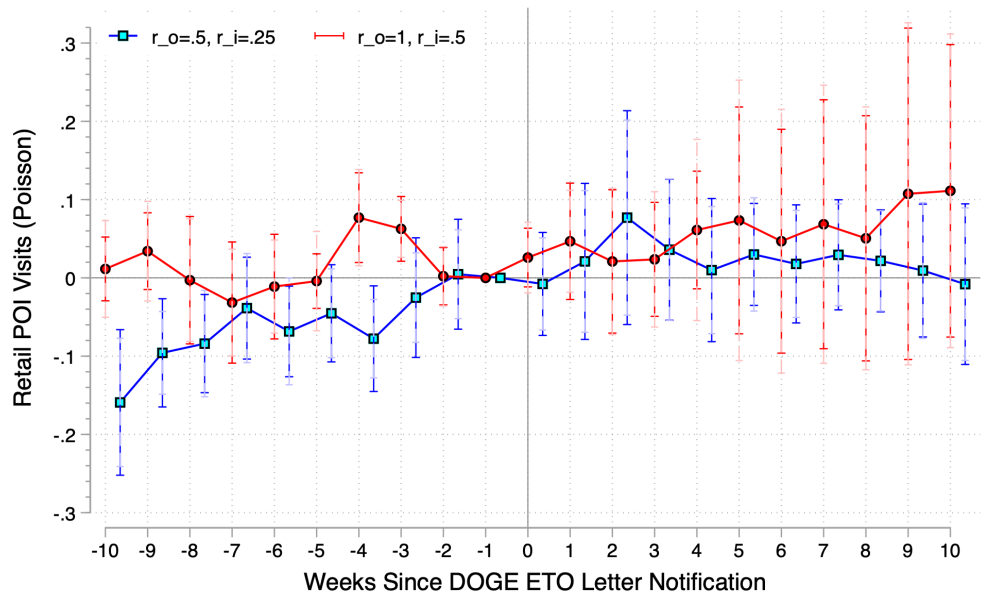
FIGURE 19. Dynamic Spatial Spillovers Diff-in-Diff Plot: SE Cluster at Bond CUSIP Level
(Y: Log CMBS Bond Price, Tranche: First-Loss Group, Sample Radius: 5-mile)



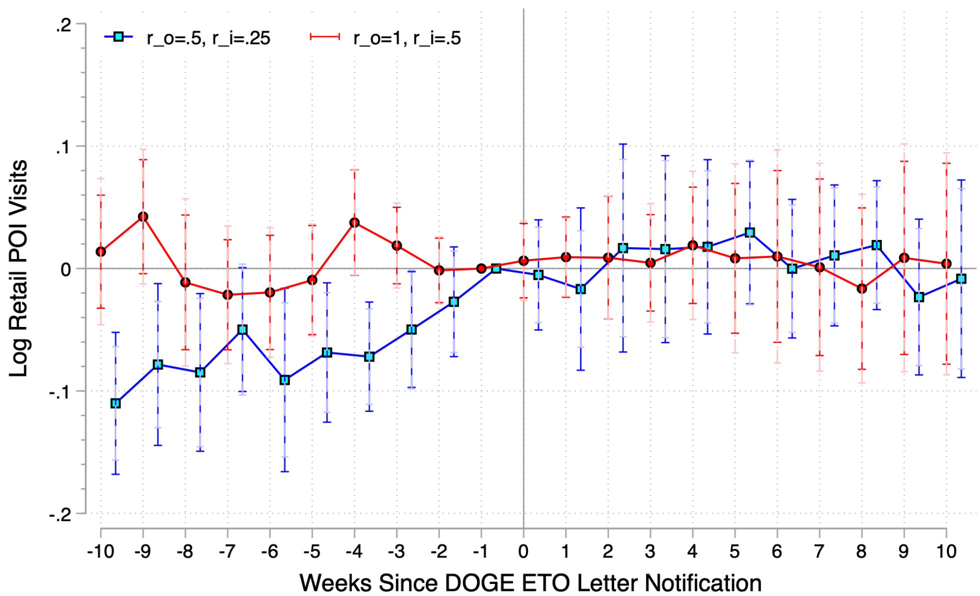
Notes: The figure plots the event study coefficients estimated from dynamic versions of equation (5.3) with log CMBS bond price as the outcome variable, applied to a 4-month symmetric window around the creation of DOGE in January 2025. We plot the estimates from four specifications: one without any fixed effects ("No FE"), then one adding in discrete distance band ring fixed effects and time fixed effects ("Ring/Time FE"), then adding CMBS deal fixed effects ("Deal"), then finally adding CUSIP fixed effects ("Bond CUSIP"). In each specification, we set $t = -1$ to be the reference period, corresponding to January 2025. This reference period choice reflects the fact that the first set of DOGE terminations for the Washington, D.C. area was sent to landlords and tenants on January 30, 2025. We restrict our sample to the first-loss group of tranches with the properties that are within the set of 5-mile radii from the properties with the canceled federal leases in Washington, D.C. We continue to follow Flynn and Ghent (2018) in defining the FLG as consisting of tranches which have a rating of CCC, or CCC+, or are unrated. We winsorize bond prices at the 1st and 99th percentiles. 95% confidence intervals obtained from clustering standard errors at the bond CUSIP level.

FIGURE 20. Null Effects of Terminated Federal Leases on Nearby Retail Foot Traffic

A. Poisson Regression Results

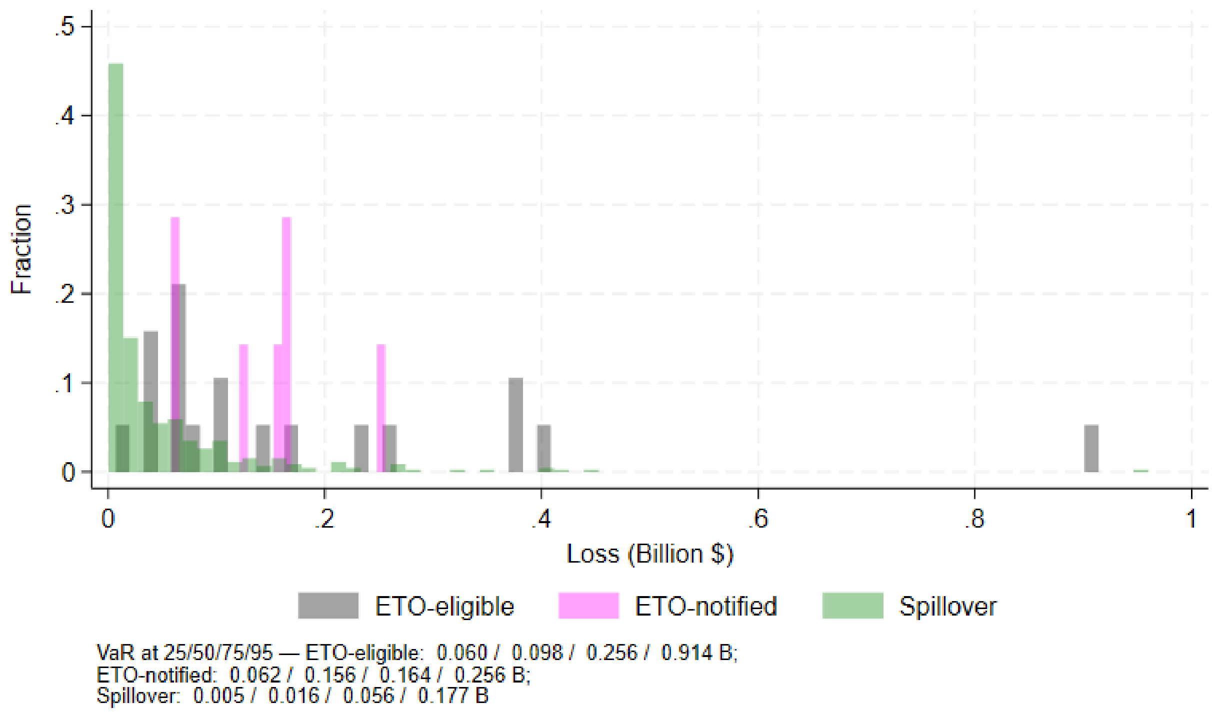


B. OLS Estimation Results



Notes: We plot event study coefficients from estimating the ring difference-in-differences specification (6.1) applied to the Advan Research (2022) foot traffic data for retail points of interest (POI). The outcome variable in all regressions is the number of visits to a POI in a given week t. In Panel A, we assume the number of visits is distributed in a Poisson fashion, following Cohn et al. (2022). In Panel B, we define the outcome as the log number of visits to a retail POI and estimate (6.1) by simple OLS. Since we restrict to a balanced panel of POIs with strictly positive foot traffic in each week, the log transform does not omit any observations. For both the Poisson and OLS approaches, we test for spillovers at different levels of proximity by varying the pair of inner and outer radii $(\mathfrak{r}_{inner}, \mathfrak{r}_{outer})$ from $(0.25, 0.5)$ miles to $(0.5, 1)$ miles. We identify POIs as retail establishments based on whether they belong to one of four two-digit NAICS sectors Retail Trade (sector codes 44–45), Arts, Entertainment, and Recreation (sector code 71) or Accommodation and Food Services (sector code 72). 95% confidence intervals in solid bars obtained from Conley (2008) standard errors with a maximal spatial correlation distance cutoff parameter determined for each set of inner and outer radii $(\mathfrak{r}_{inner}, \mathfrak{r}_{outer})$. 95% confidence intervals in lighter dashed bars obtained from clustering standard errors at the Census block group level.

FIGURE 21. Simulated Loss Distributions
ETO-Eligible (Non-Exercised) vs. ETO-Notified (Exercised) vs. Private Tenant



Notes: The figure plots the simulated loss distributions under early termination option (ETO) risk for three groups of properties: those that are eligible for ETO exercise but not notified (gray), those that have received formal ETO notifications (pink), and private-lease properties within 5-mile radii indirectly affected through spatial spillovers (green). We define ETO-eligible leases as those which are already in the soft term of the lease as of January 2025 when DOGE was created. Losses are measured in billions of dollars, based on Monte Carlo simulations of property value declines over a five-year horizon using the jump-diffusion processes in (E.1)–(E.3) calibrated to observed termination rates and hedonic value baselines. We summarize the Value at Risk (VaR) for the $1 - \alpha \in \{25\%, 50\%, 75\%, 95\%\}$ levels. For instance, at the 95% Value at Risk (VaR) losses exceed $914 million, $256 million, and $177 million in the worst 5% of scenarios for ETO-eligible, ETO-notified, and private tenant properties, respectively. See Appendix E for details.

Tables

TABLE 1. Telework Policies of Federal Agencies with Securitized GSA Leases in the Washington, D.C. Metro Area (DOGE-Notified & ETO-Eligible Leases)

Status	Agency	Days/Week In-Office (est.)
DOGE-notified	Department of Veterans Affairs (VA)	2.5
DOGE-notified	Federal Energy Regulatory Commission (FERC)	2
DOGE-notified	Department of Homeland Security (DHS)	3 (est.)
DOGE-notified	General Services Administration (GSA)	1-2 (est.)
DOGE-notified	Federal Emergency Management Agency (FEMA)	2
DOGE-notified	Department of Housing and Urban Development (HUD)	1
DOGE-notified	Department of the Treasury	2.5
DOGE-notified	Internal Revenue Service (IRS)	2.5
DOGE-notified	Department of Energy (DOE)	3
DOGE-notified	Federal Aviation Administration (FAA)	2
ETO-eligible	GSA National Capital Region 11	2 (est.)
ETO-eligible	U.S. Navy	3
ETO-eligible	Department of Housing and Urban Development (HUD)	1
ETO-eligible	U.S. Federal Labor Relations Authority	3 (est.)
ETO-eligible	The Public Defender Service	3 (est.)
ETO-eligible	U.S. Chemical Safety Board	3 (est.)
ETO-eligible	United States Postal Service (USPS)	3
ETO-eligible	AmeriCorps	3 (est.)
ETO-eligible	Federal Mediation and Conciliation Service	4 (est.)
ETO-eligible	U.S. Office of Government Ethics	1-2 (est.)
ETO-eligible	National Aeronautics and Space Administration (NASA)	2.5 (est.)
ETO-eligible	NASA Office of Inspector General	3 (est.)
ETO-eligible	Office of the Comptroller of the Currency (OCC)	3 (est.)
ETO-eligible	Federal Housing Finance Agency (FHFA)	1-2 (est.)
ETO-eligible	Federal Trade Commission (FTC)	3 (est.)
ETO-eligible	National Endowment for the Humanities (NEH)	1-2 (est.)
ETO-eligible	National Endowment for the Arts (NEA)	2.5 (est.)
ETO-eligible	National Institute for Occupational Safety and Health (NIOSH)	2.5 (est.)
ETO-eligible	Networking and Information Technology Research and Development (NITRD)	2.5 (est.)
ETO-eligible	National Transportation Safety Board (NTSB)	1-2 (est.)
ETO-eligible	Argonne National Laboratory	2.5 (est.)
ETO-eligible	Court Services and Offender Supervision Agency (CSOSA)	1-2 (est.)
ETO-eligible	Pretrial Services Agency for the District of Columbia (PSA)	2.5 (est.)
ETO-eligible	Public Defender Service for the District of Columbia (PDS)	2.5 (est.)
ETO-eligible	Federal Retirement Thrift Investment Board (FRTIB)	3 (est.)
ETO-eligible	U.S. Agency for International Development (USAID)	3
ETO-eligible	National Science Foundation (NSF)	2

Notes: The table reports average in-office days per week for federal agency employees attached to DOGE-notified and ETO-eligible (but not notified by DOGE) General Services Administration (GSA) leases in the Washington, D.C. metro area. We define ETO-eligible leases as those which will be in the soft term of the lease at any point during the current presidential administration, or between January 2025 and January 2029. We extract property addresses for DOGE-notified and ETO-eligible GSA leases in the Trepp data, which consists of properties with CMBS loans (see Section 4 for details). We identify the federal agency tenants located at Trepp GSA addresses from CoStar. We classify by hand estimated average days per week worked from home for each agency from the Appendix of the 2024 OMB Report to Congress on Telework and Real Property Utilization (Office of Management and Budget, 2024) If in-office days per week are not explicitly reported, we infer from the telework participation numbers from the OMB report and estimate the average days per week worked in the office; we denote such cases as ("est.").

TABLE 2. Summary Statistics
(CMBS Bond Prices, NOI)

Panel (A)	Full Sample
Avg Log (CMBS Bond Price)	3.74
Std Dev. Log (CMBS Bond Price)	1.62
Observations	1771195
Avg Log (Net Operating Income)	15.04
Std Dev. Log (Net Operating Income)	1.98
Observations	1544494

Panel (B)	ETO Exercisable, Notification Not Sent			
	First Loss Group	Mezzanine	Senior	Unclassified
Avg Log (CMBS Bond Price)	3.28	4.09	3.25	3.53
Std Dev. Log (CMBS Bond Price)	1.66	0.86	2.34	1.81
Observations	3127	42822	28422	1407
Avg Log (Net Operating Income)	15.67	16.70	16.31	14.51
Std Dev. Log (Net Operating Income)	2.00	1.33	1.29	1.73
Observations	1496	6756	4814	566

Panel (C)	ETO Exercisable, Notification Sent			
	First Loss Group	Mezzanine	Senior	Unclassified
Avg Log (CMBS Bond Price)	2.98	3.92	3.42	4.16
Std Dev. Log (CMBS Bond Price)	1.26	1.23	2.15	1.08
Observations	1496	6827	5114	533
Avg Log (Net Operating Income)	16.72	16.29	16.17	16.52
Std Dev. Log (Net Operating Income)	0.15	0.48	0.46	0.40
Observations	1496	6756	4814	566

Notes: The table reports summary statistics in our nationwide Trepp CMBS sample for the mean, standard deviation, and number of observations for the log of CMBS bond price and NOI. Panel (A) documents the full sample of CMBS bonds matched to Trepp properties for 2020 onward. Panel (B) and (C) are stratified by the government cancellation notification sent or not respectively within the ETO exercisable leases. Each column corresponds to the tranche group. We follow Flynn and Ghent (2018) in defining the tranche groups according to their bond ratings. The First Loss Group (FLG) as consists of tranches which have a rating of CCC, or CCC+. We classify mezzanine tranches as those rated below AAA but above CCC+. Senior tranches are those with a AAA rating. Unclassified bonds are those which are "unrated" but do not have a missing value for the rating provided by the rating agencies. In cases where the bond receives multiple agency ratings, we use the S&P rating. If the S&P rating is unavailable, we use the Fitch rating. Finally, if both the S&P and Fitch ratings are unavailable, we adopt the Moody's rating.

TABLE 3. Difference-in-Differences Regressions (Y: Log CMBS Bond Prices)

	(1)	(2)	(3)	(4)
Post	0.025	0.022		
	(0.017)	(0.014)		
DOGE	-0.438	-0.336*	-0.336*	
	(0.290)	(0.175)	(0.176)	
$DOGE \times Post$	-0.040**	-0.037**	-0.037**	-0.024**
	(0.018)	(0.015)	(0.015)	(0.010)
Adj-R^2	0.025	0.220	0.213	0.999
Observations	742	742	742	742
Property Zip FE		✓	✓	
Bond Time FE			✓	✓
Bond CUSIP FE				✓

Notes: The table reports our pooled difference-in-differences regression results with log of CMBS bond prices as the outcome variable, estimated according to equation (5.1). *Post* equals 1 if a period is after January 20, 2025, *DOGE* equals 1 if the U.S. government sent ETO notifications. The sample consists of the leases that are in the first loss group tranche, with not-yet ETO-eligible leases serving as the control group. We define not-yet ETO-eligible leases as those with a termination right date that falls under the current presidential administration, or from January 2025 to January 2029. The property zip location, deal distribution time, bond tranche, and bond CUSIP fixed effects are cumulatively added in Columns (1), (2), (3), and (4) respectively. Standard errors clustered by bond CUSIP in parentheses. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

TABLE 4. Difference-in-Differences Regressions (Y: Log NOI)

	(1)	(2)	(3)	(4)
Post	0.014	-0.003		
	(0.015)	(0.003)		
DOGE	-0.062	-1.351***	-1.351***	
	(0.272)	(0.041)	(0.041)	
DOGE × *Post*	-0.068***	-0.051***	-0.051***	-0.054***
	(0.019)	(0.013)	(0.013)	(0.013)
Adj-R^2	-0.003	0.986	0.986	1.000
Observations	742	742	742	742
Property Zip FE		✓	✓	
Bond Time FE			✓	✓
Bond CUSIP FE				✓

Notes: The table reports our pooled difference-in-differences regression results with log of net operating income (NOI) as the outcome variable, estimated according to equation (5.1). *Post* equals 1 if a period is after January 20, 2025, *DOGE* equals 1 if the U.S. government sent ETO notifications. The sample consists of the leases that are in the first loss group tranche, with not-yet ETO-eligible leases serving as the control group. We define not-yet ETO-eligible leases as those with a termination right date that falls under the current presidential administration, or from January 2025 to January 2029. The property zip location, deal distribution time, bond tranche, and bond CUSIP fixed effects are cumulatively added in Columns (1), (2), (3), and (4) respectively. Standard errors clustered by bond CUSIP in parentheses. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

TABLE 5. Difference-in-Differences Regressions (Y: Log Avg Weighted DSCR)

	(1)	(2)	(3)	(4)
Post	0.016***	0.016***		
	(0.003)	(0.003)		
DOGE	0.313***	0.356***	0.356***	
	(0.068)	(0.083)	(0.084)	
$DOGE \times Post$	-0.027***	-0.027***	-0.027***	-0.025***
	(0.004)	(0.004)	(0.004)	(0.003)
Adj-R^2	0.134	0.183	0.176	0.999
Observations	742	742	742	742
Property Zip FE		✓	✓	
Bond Time FE			✓	✓
Bond CUSIP FE				✓

Notes: The table reports our pooled difference-in-differences regression results with log of weighted average debt service coverage ratio as the outcome variable, estimated according to equation (5.1). *Post* equals 1 if a period is after January 20, 2025, *DOGE* equals 1 if the U.S. government sent ETO notifications. The sample consists of the leases that are in the first loss group tranche, with not-yet ETO-eligible leases serving as the control group. We define not-yet ETO-eligible leases as those with a termination right date that falls under the current presidential administration, or from January 2025 to January 2029. The property zip location, deal distribution time, bond tranche, and bond CUSIP fixed effects are cumulatively added in Columns (1), (2), (3), and (4) respectively. Standard errors clustered by bond CUSIP in parentheses. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

TABLE 6. Spatial Difference-in-Differences Regressions (Y: Log CMBS Bond Prices)

	(1)	(2)	(3)	(4)
Post	-0.111*			
	(0.060)			
Spillover	-0.139	-0.172	0.051***	0.041**
	(0.129)	(0.131)	(0.019)	(0.016)
Spillover × *Post*	-0.149**	-0.136**	-0.108***	-0.088**
	(0.066)	(0.061)	(0.040)	(0.035)
Adj-R^2	0.004	0.007	0.374	0.978
Observations	8015	8015	8014	8008
Ring-Time FE		✓	✓	✓
Deal FE			✓	
Bond CUSIP				✓

Notes: The table reports our spatial difference-in-differences regression results with log of CMBS bond prices as the outcome variable, estimated according to equation (5.3). *Post* equals 1 if a period is after January 20, 2025, *Spillover* equals 1 if the bond-deal involves a private tenant and the underlying space being rented is located within a 5-mile radius of an ETO-exercised lease. The sample consists of the leases that are within a 5-mile radius of the DOGE-led canceled leases and first loss group tranche. The ring-time, deal, and bond CUSIP fixed effects are cumulatively added in Columns (2), (3), and (4) respectively. Ring fixed effects refer to an indicator equal to one if the bond-deal (i, c) lies within the 1-mile radius of DOGE-canceled lease r. Standard errors clustered by bond CUSIP in parentheses. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

TABLE 7. Spatial Triple Difference-in-Differences Regressions (Y: Log CMBS Bond Prices)

	(1)	(2)	(3)	(4)
Post × *Private*	-0.189**	-0.185**	-0.173***	-0.069**
	(0.091)	(0.091)	(0.050)	(0.029)
Post × *Ring*	-0.031	-0.025	-0.023	-0.037
	(0.024)	(0.026)	(0.026)	(0.023)
Ring × *Private* × *Post*	-0.067	-0.073	-0.050	-0.043
	(0.058)	(0.059)	(0.046)	(0.042)
Adj-R^2	0.005	0.008	0.374	0.978
Observations	8015	8015	8014	8008
Ring-Time FE		✓	✓	✓
Deal FE			✓	
Bond CUSIP				✓

Notes: The table reports our spatial difference-in-differences regression results with log of CMBS bond prices as the outcome variable, estimated according to equation (5.4). *Post* equals 1 if a period is after January 20, 2025, *Private* equals 1 if the property has a private tenant (i.e., non-federal lease), and *Ring* equals 1 if the lease is within the 1-mile radii of the ETO-canceled federal lease buildings in Washington D.C. The sample consists of the leases that are within a 5-mile radius of the DOGE-led canceled leases and first loss group tranche. The ring and time, deal, and bond CUSIP fixed effects are cumulatively added in Columns (2), (3), and (4) respectively. Standard errors clustered by bond CUSIP in parentheses. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

TABLE 8. Summary of Simulated Value at Risk (VaR) and Expected Shortfall (ES)
Washington, D.C. Office Property Value Losses ($ Billions)

	Value at Risk				Expected Shortfall			
$1 - \alpha =$	25%	50%	75%	95%	25%	50%	75%	95%
ETO-eligible	0.060	0.098	0.256	0.914	0.239	0.307	0.464	0.914
ETO-notified	0.062	0.156	0.164	0.256	0.154	0.185	0.256	0.256
Spillover	0.005	0.016	0.056	0.177	0.060	0.084	0.137	0.298
Total losses	0.127	0.270	0.476	1.347	0.453	0.576	0.857	1.468

Notes: The table summarizes the Value at Risk (VaR) and expected shortfall (ES) notions of office property valuation losses from the DOGE federal lease cancellations at various levels of the α% worst outcomes. We report losses in billions of dollars. Following the definitions used throughout the paper, ETO-eligible refers here to losses incurred by properties which are already eligible for early termination as of January 2025, ETO-notified refers to losses among properties with leases canceled by DOGE, and spillover refers to non-GSA properties within a 5-mile radius of those actually canceled. We follow the simulation procedures described in Section 7, under the calibration outlined in Appendix E.

Online Appendix to

Pricing Government Contract Risk Premia: Evidence from the 2025 Federal Lease Terminations

Soon Hyeok Choi (Rochester Institute of Technology) & Cameron LaPoint (Yale SOM)

A Model Extensions

A.1 Insurance Premia

Given losses stemming from the mispricing of the government contract risk premium associated with an early termination option (ETO), a natural question arises: if a market for ETO contingencies existed in the form of insurance what would be the price? To address this, we derive a closed-form solution for the insurance premium corresponding to an ETO contingency.

A federal lease is typically divided into a firm term and a soft term, each comprising roughly half of the contract term. As the firm term carries no credit risk due to its non-cancelable nature, our analysis focuses on the soft term. Suppose the soft term spans a fixed horizon, $[t_0, t_m]$ with $m \in \mathbb{N}$. The contract obligates the federal tenant to pay the base rent R from t_1 to t_m. Since the federal tenant holds a long position in the ETO (put option), it pays a regular insurance premium (c) to the landlord during the soft term. However, the federal tenant can exercise its ETO during any time in this soft term. The stopping time, τ, represents the period that the tenant officially notifies its intent to terminate the lease early, where $\tau \in (t_{h-1}, t_h]$ for $h \in \{1, ..., m\}$.

The advance notification period, denoted by α, serves as a grace period during which the landlord is formally informed of the tenant's intention to exercise the ETO. This notification initiates a transitional window allowing the landlord to commence re-leasing efforts in anticipation of the forthcoming vacancy, thereby partially mitigating the risk of rental income disruption. Throughout the interval $[t_h, \lceil\tau\rceil + \alpha]$, the federal tenant remains contractually obligated to continue rental payments. Upon the expiration of this notice period, starting at $\lceil\tau\rceil + \alpha + 1$, the financial responsibility for the vacant space shifts to the landlord, who must then absorb the rental cost until a replacement tenant is secured. Since the time of successful re-leasing is uncertain, $\eta \in (t_{k-1}, t_k]$ for $k > \lceil\tau\rceil + \alpha + 1$ denotes a stopping time marking the end of the vacancy period, beyond which the landlord ceases rent payments beginning at t_k. Figure 2 summarizes the timeline of key events with an exercised ETO.

Proposition 6. *Suppose there is one landlord and one federal tenant. Consider a soft term in which the tenant can exercise an early termination option (ETO). Suppose the tenant sends its ETO notification at τ with the grace period α and the random time η at which a replacement tenant can occur. Then the ETO* *insurance premium is:*

$$c = \frac{RE_{\mathbb{Q}}\left[\sum_{h=\lceil\tau\rceil+\alpha+1}^{\lceil\eta\rceil\wedge t_m} \mathbb{1}_{\{t_0<\tau\leq t_m\}} e^{-\int_0^{t_h} r_u du}\right]}{E_{\mathbb{Q}}\left[\sum_{k=1}^{\lceil\tau\rceil+\alpha} p(0,t_k)\right]} \tag{A.1}$$

Proof. See Appendix B.6 for the derivations. □

B Proofs

B.1 Proof of Proposition 1

If the rent $\widetilde{R}$ correctly priced the government contract risk premium associated with the ETO:

$$\frac{V(t)}{B(t)} = \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{\tau} \frac{\widetilde{R}}{B(s)} + \sum_{s=\eta+1}^{T} \frac{\widetilde{R}}{B(s)} + \frac{V(T)}{B(T)}\right] \tag{B.1}$$

$$\implies \frac{V(t)}{B(t)} - \mathbb{E}_{\mathbb{Q}}\left[\frac{V(T)}{B(T)}\right] = \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{\tau} \frac{\widetilde{R}}{B(s)} + \sum_{s=\eta+1}^{T} \frac{\widetilde{R}}{B(s)}\right] \tag{B.2}$$

$$\implies \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{T} \frac{R}{B(s)}\right] = R\,\mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{\tau} \frac{1}{B(s)} + \sum_{s=\eta+1}^{T} \frac{1}{B(s)}\right] \tag{B.3}$$

$$\implies R\,\mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{\tau} \frac{1}{B(s)} + \sum_{s=\tau+1}^{\eta} \frac{1}{B(s)} + \sum_{s=\eta+1}^{T} \frac{1}{B(s)}\right] = \widetilde{R}\,\mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{\tau} \frac{1}{B(s)} + \sum_{s=\eta+1}^{T} \frac{1}{B(s)}\right] \tag{B.4}$$

$$\implies \widetilde{R} = R\,\mathbb{E}_{\mathbb{Q}}\left[1 + \frac{\sum_{s=t+1}^{\tau} \frac{1}{B(s)}}{\sum_{s=t+1}^{\tau} \frac{1}{B(s)} + \sum_{s=\eta+1}^{T} \frac{1}{B(s)}}\right] > R \tag{B.5}$$

□

B.2 Proof of Proposition 2

$$L(t) = \frac{V(t) - \widetilde{V}(t)}{B(t)} = \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=\tau+1}^{\eta} \frac{R}{B(s)}\right] \tag{B.6}$$

$$= R \sum_{s=t+1}^{T} p(t,s)\,\mathbb{Q}(\tau \leq s)\,\mathbb{Q}(s \leq \eta). \tag{B.7}$$

$$\text{Note that } \mathbb{Q}(\tau \leq s)\,\mathbb{Q}(s \leq \eta) = \left(1 - e^{-\lambda_\tau s}\right)\left(e^{-\lambda_\eta s}\right) = e^{-\lambda_\eta s} - e^{-(\lambda_\tau + \lambda_\eta)s} \tag{B.8}$$

$$= R \sum_{s=t+1}^{T} p(t,s)\left[e^{-\lambda_\eta s} - e^{-(\lambda_\tau+\lambda_\eta)s}\right] \tag{B.9}$$

□

B.3 Proof of Proposition 3

Using the independence termination and interest rates, noting that $\mathbb{E}_{\mathbb{Q}}\left[\mathbb{1}_{\tau>s}\right] = e^{-\lambda_\tau(s-t)}$ yields

$$\phi(t) = N \cdot \mathbb{E}_{\mathbb{Q}}\left[\sum_{s=t+1}^{T} \frac{\theta_0\widetilde{R} + \theta_1 R \mathbb{1}_{\tau>s}}{B(s)} + \frac{\theta_0\widetilde{V}(T) + \theta_1 V(T)\mathbb{1}_{\tau>T}}{B(T)}\right] \tag{B.10}$$

$$= N\left[\sum_{s=t+1}^{T} \mathbb{E}_{\mathbb{Q}}\left[\frac{B(t)}{B(s)}\right]\left[\theta_0\widetilde{R} + \theta_1 R e^{-\lambda_\tau(s-t)}\right] + \mathbb{E}_{\mathbb{Q}}\left[\frac{B(t)}{B(T)}\right]\left[\theta_0\widetilde{V}(T) + \theta_1 V(T) e^{-\lambda_\tau(T-t)}\right]\right]$$

$$= N\left[\sum_{s=t+1}^{T} p(t,s)\left[\theta_0\widetilde{R} + \theta_1 R e^{-\lambda_\tau(s-t)}\right] + p(t,T)\left[\theta_1 V(T) + \theta_1 V(T) e^{-\lambda_\tau(T-t)}\right]\right] \tag{B.11}$$

Hence,

$$\frac{\partial\phi(t)}{\partial\lambda_\tau} = -N\left\{\sum p(t,s)_{s=t+1}^{T}\left[\theta_1 R(s-t)e^{-\lambda_\tau(s-t)}\right] - p(t,T)\left[\theta_1 V(T)(T-t)e^{-\lambda_\tau(T-t)}\right]\right\} < 0$$

□

B.4 Proof of Proposition 4

Since $\widehat{\phi}(t,i)$ is linear in $\widehat{\pi}(s,i)$, it suffices to show the marginal effects of the state transition probabilities and intensities on $\widehat{\pi}(s,i)$. First, increasing q_{LH} makes transitions from L (baseline regime) to H (exit regime) more likely. Since $\lambda_\tau^H > \lambda_\tau^L$, spending more time in the exit regime leads to a higher effective hazard rate. Hence, increasing q_{LH} increases the rate at which the survival probability decays:

$$\frac{\partial \widehat{\pi}(s,i)}{\partial q_{LH}} < 0 \implies \frac{\partial \widehat{\phi}(t,i)}{\partial q_{LH}} < 0$$

By the same intuition, spending more time in the baseline regime slows down the decay of survival probability:

$$\frac{\partial \widehat{\pi}(s,i)}{\partial q_{HL}} > 0 \implies \frac{\partial \widehat{\phi}(t,i)}{\partial q_{HL}} > 0$$

Regardless of the regime $i \in \{L, H\}$, increasing λ_τ^i directly increases the hazard rate in the exit regime without affecting the regime transition probability matrix Q. Whenever the system is in the exit regime, survival decays faster:

$$\frac{\partial \widehat{\pi}(s,i)}{\partial \lambda_\tau^i} < 0 \implies \frac{\partial \widehat{\phi}(t,i)}{\partial \lambda_\tau^i} < 0$$

□

B.5 Proof of Proposition 5

$$\frac{\partial \widehat{\phi}_c(t,i)}{\partial \varsigma} = N\left[\sum_{s=t+1}^{T} p(t,s)\left[-\theta_0 \bar{R}(1-\pi(s,i))\right] + p(t,T)\left[-\theta_0 \bar{V}(T)(1-\pi(T,i))\right]\right] < 0 \tag{B.12}$$

$$\frac{\partial \widehat{\phi}_c(t,i)}{\partial \Xi(s,i)} = \begin{cases} N \cdot p(t,s)\theta_0 \bar{R} > 0, & s < T \\ N \cdot p(t,T)\theta_0 \bar{R} > 0, & s = T \end{cases} \tag{B.13}$$

□

B.6 Proof of Proposition 6

Using the same argument as above:

$$V(0) = 0 \quad \text{given that the time } t = 0 \text{ fair value of the ETO contract is zero.}$$

$$0 = E_{\mathbb{Q}}\Big[\sum_{k=1}^{\lceil \tau \rceil + \alpha} c e^{-\int_0^{t_k} r_u du} - \sum_{h=\lceil \tau \rceil + \alpha + 1}^{\lceil \eta \rceil \wedge t_m} R \mathbb{1}_{\{t_0 < \tau \leq t_h\}} e^{-\int_0^{t_j} r_u du}\Big] \tag{B.14}$$

$$\implies c = \frac{RE_{\mathbb{Q}}\left[\sum_{h=\lceil\tau\rceil+\alpha+1}^{\lceil\eta\rceil\wedge t_m} \mathbb{1}_{\{t_0<\tau\leq t_m\}} e^{-\int_0^{t_h} r_u du}\right]}{E_{\mathbb{Q}}\left[\sum_{k=1}^{\lceil\tau\rceil+\alpha} p(0,t_k)\right]} \tag{B.15}$$

□

C Additional Results for Foot Traffic Analysis

This appendix presents supplemental information about the Advan Research data on foot traffic and robustness checks corresponding to the tests for consumption externalities from Section 6.4.

C.1 Foot Traffic Sample Creation

We obtain data on foot traffic for points of interests (POIs) from Advan Research (2022). The data tracks the number of visits to POIs using cellphone pings from an anonymized panel of mobile devices, allowing us to observe weekly volume of visits and visitors. Our sample consists of a balanced panel of POIs for all weeks between June 2023 and June 2025. POIs in our data are *static* in the sense that our set of POIs and their characteristics are fixed as of June 2023, and we cannot add or drop POIs based on unobserved establishment turnover or openings. Instead, we exclude POIs that experience zero visits in any week of our sample period to avoid tracking POIs that experience business closure. Our main analysis further subsets to POIs that we identify as retail establishments based on belonging to one of four two-digit NAICS sectors *Retail Trade* (sector codes 44–45), *Arts, Entertainment, and Recreation* (sector code 71) or *Accommodation and Food Services* (sector code 72). We present robustness analyses on a non-retail sample that includes all other NAICS sectors in Appendix C.2.

A common concern with foot traffic data inferred from cellphone locations is that the underlying panel of mobile devices might not uniformly cover the area of interest (Hou et al., 2025). Given that we exclusively focus on the well-populated area of D.C., we expect this to be less of an issue in our application. However, our focus on a static panel of POIs with always non-zero foot traffic might induce selective coverage of POIs. For example, POIs further away from the more frequented areas in the center might experience visits less regularly and therefore could be more prone to exclusion from the sample. Figure C.1 maps out the share of POIs that are excluded from the sample due to having zero weekly foot traffic. Census tracts located further away from the central business district (CBD) are subject to a higher share of excluded POIs. Fortunately, 13 out of the 15 locations with DOGE terminated leases are located close to the CBD where POI coverage is the greatest. In addition, our spillover design limits variation in POI coverage by only comparing POIs in small concentric circles around DOGE terminated leases, among which POI coverage appears to be relatively stable. We exclude two more canceled leases which are located in high coverage rate areas but were canceled later in the sample after investors would have witnessed rescission decisions.

C.2 Robustness: Foot Traffic Responses of Non-Retail Establishments

As a sanity check, we re-estimate our specifications using the foot traffic data but for the sample of non-retail leases, defined by the complement set of any of the 2-digit NAICS codes not classified as retail. We use this non-retail definition rather than further defining office and industrial groups, as there is no clear way to map NAICS codes for the establishment into a use of the property. We again estimate the difference-in-differences specification in equation (6.1) by Poisson regression and OLS. We continue to find a statistically flat pattern of foot traffic around DOGE lease cancellation events. If anything, foot traffic to non-retail establishments slightly increases in the month after termination announcements.

C.3 Robustness: Early vs. Late Ring Difference-in-Differences

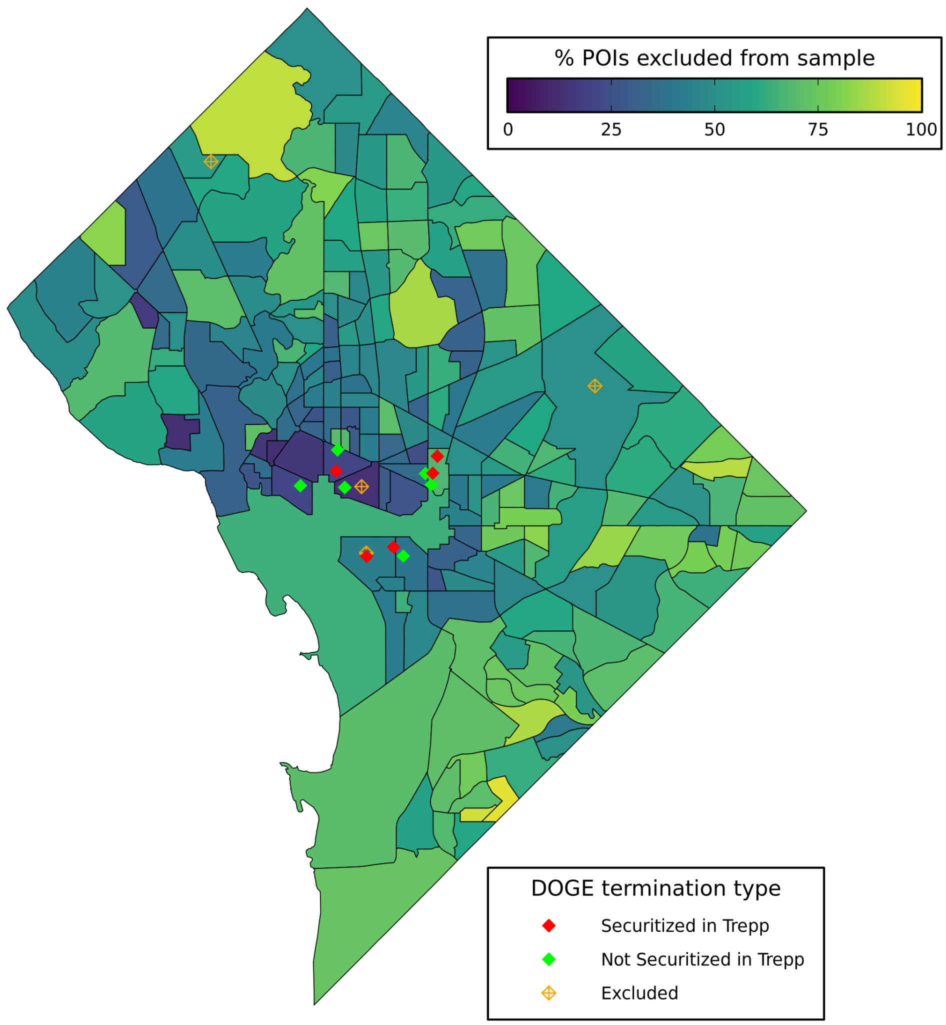
An alternative approach to estimating spillovers on foot traffic makes use of the staggered termination of leases, comparing the evolution of foot traffic around federal leases that DOGE terminated first to the evolution of foot traffic around federal leases terminated later. We can implement this research design with the Advan data which are at a weekly frequency, unlike our data on CMBS prices and property performance which are only available monthly and therefore do not divide up leases into clear treatment cohorts.

To implement this early vs. late design, we estimate the following regression specification:

$$Y_{j,r,s,t} = \sum_{t=-5,e\neq-1}^{+4} \beta_t \cdot Spillover_{i,t} + \mu_i + \delta_{s,t} + \epsilon_{j,r,s,t} \tag{C.1}$$

Quantities in this equation are defined as in equation (6.1), except for $Spillover_{i,t}$ which now takes value one if the POI is located within 0.5 miles or 0.25 miles of a terminated DOGE lease and if the relative time to the termination of that DOGE lease equals t. We define the control group as those POIs which are close to a DOGE lease but which have not been terminated as of period t and will be in the future. Appendix Figure C.3 gives a visual representation of early and late-treated POIs. We exclude from the estimation rings defined by two leases canceled after March 2025, as such rings may differ from earlier cohorts due to the informational content of being listed on the DOGE website. For example, landlords for leases that were newly canceled in May 2025 may have already incorporated signals received from the rescission of leases between March and May, resulting in little change in the status quo of business patterns after the actual termination announcement. Our results are less precise but qualitatively similar – meaning we still find null effects on foot traffic – if we include these later cohorts of rings in the sample.

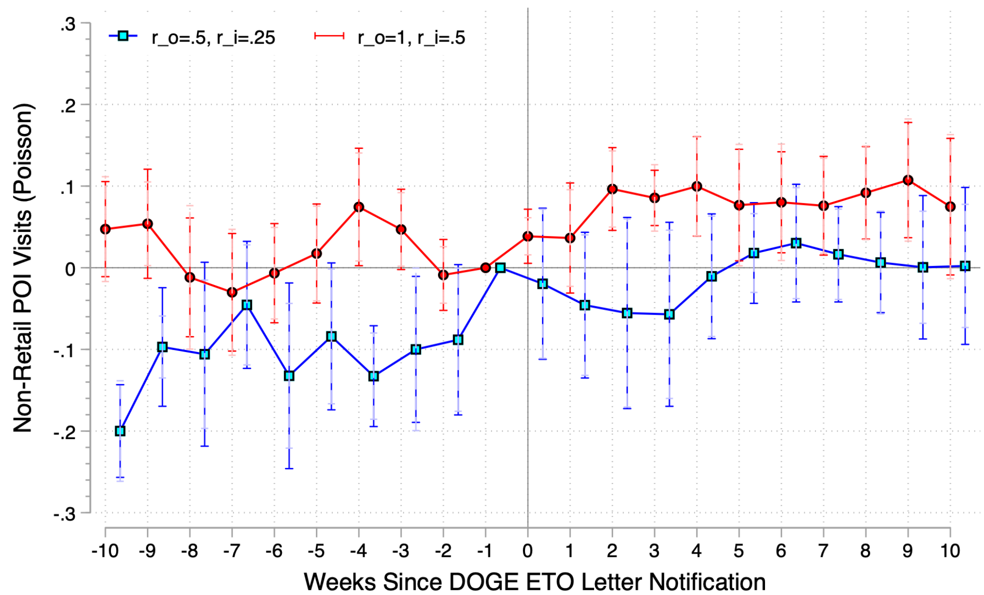
FIGURE C.1. Spatial Distribution of Points of Interest (POIs) Excluded from Advan Sample



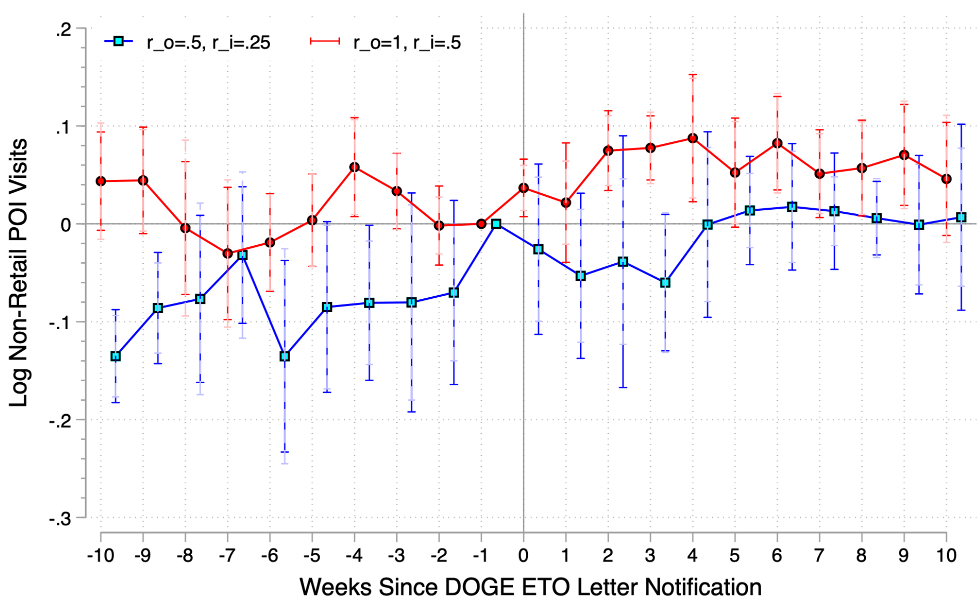
Notes: The map plots the percentage of points of interest (POIs) that are excluded from the Advan foot traffic analysis sample due to having zero foot traffic in at least one month during the sample period. Each polygon represents a census tract in Washington, D.C., with the color scale indicating the percentage of POIs located in the census tract that were excluded from the analysis sample. Red diamonds indicate the location of DOGE terminated federal leases that are included in our Trepp analysis sample, while the green diamonds indicate the non-securitized canceled leases which are not in Trepp. The two orange kites indicate the location of DOGE-terminated federal leases that we exclude from our analysis due to their location in areas with weaker coverage.

FIGURE C.2. Null Effects of Terminated Federal Leases on Non-Retail Foot Traffic

A. Poisson Regression Results

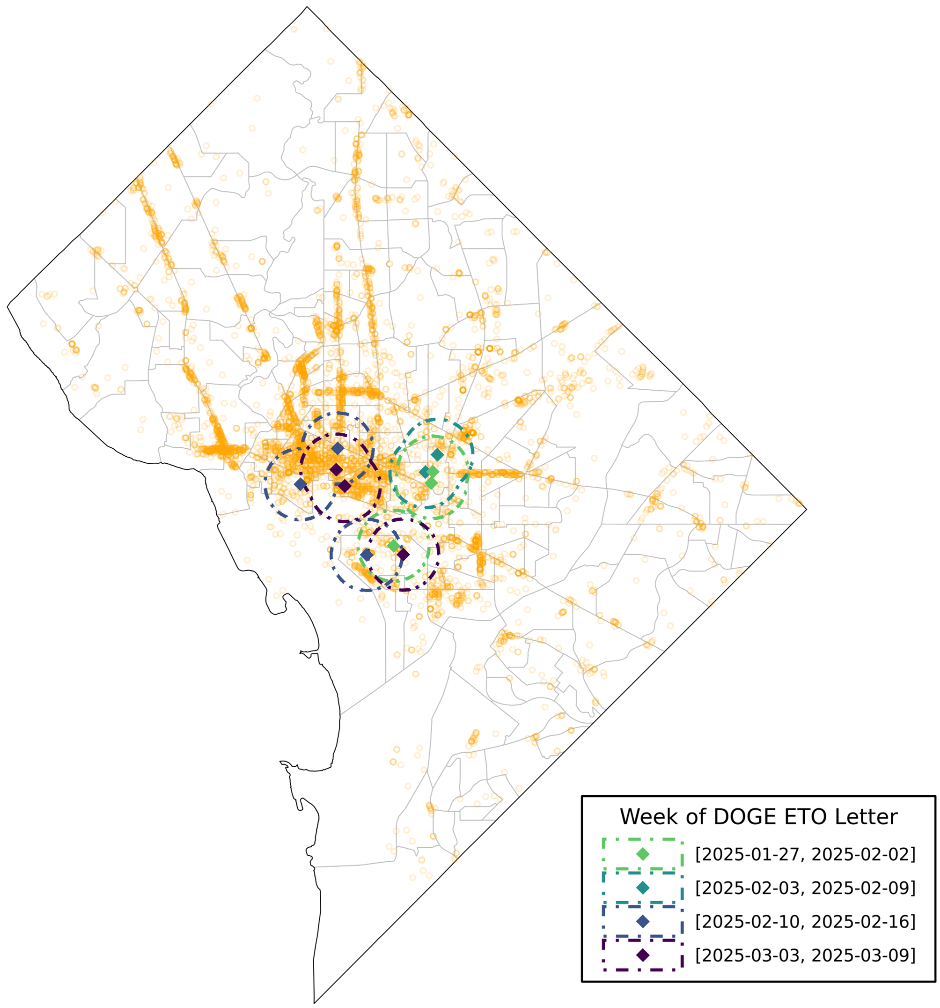


B. OLS Estimation Results



Notes: We plot event study coefficients from estimating the ring difference-in-differences specification (6.1) applied to the Advan Research (2022) foot traffic data for non-retail points of interest (POI). The outcome variable in all regressions is the number of visits to a POI in a given week t. In Panel A, we assume the number of visits is distributed in a Poisson fashion, following Cohn et al. (2022). In Panel B, we define the outcome as the log number of visits to a non-retail POI and estimate (6.1) by simple OLS. Since we restrict to a balanced panel of POIs with strictly positive foot traffic in each week, the log transform does not omit any observations. For both the Poisson and OLS approaches, we test for spillovers at different levels of proximity by varying the pair of inner and outer radii $(\mathfrak{r}_{inner}, \mathfrak{r}_{outer})$ from $(0.25, 0.5)$ miles to $(0.5, 1)$ miles. We identify POIs as non-retail establishments based on whether they do not belong to one of four retail two-digit NAICS sectors *Retail Trade* (sector codes 44–45), *Arts, Entertainment, and Recreation* (sector code 71) or *Accommodation and Food Services* (sector code 72). 95% confidence intervals in solid bars obtained from Conley (2008) standard errors with a maximal spatial correlation distance cutoff parameter determined for each set of inner and outer radii $(\mathfrak{r}_{inner}, \mathfrak{r}_{outer})$. 95% confidence intervals in lighter dashed bars obtained from clustering standard errors at the Census block group level.

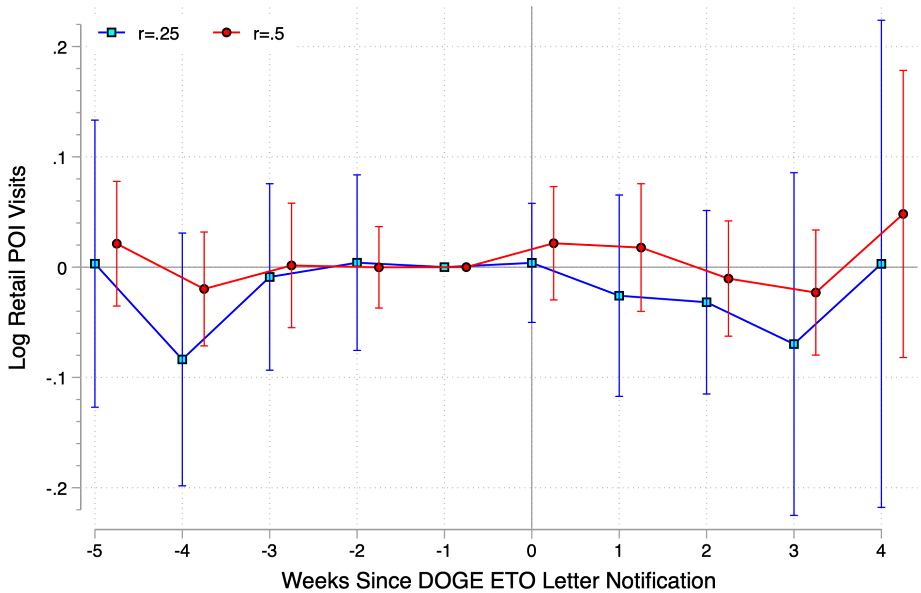
FIGURE C.3. Early and Late-Treated Rings around Terminated D.C. Federal Lease Properties



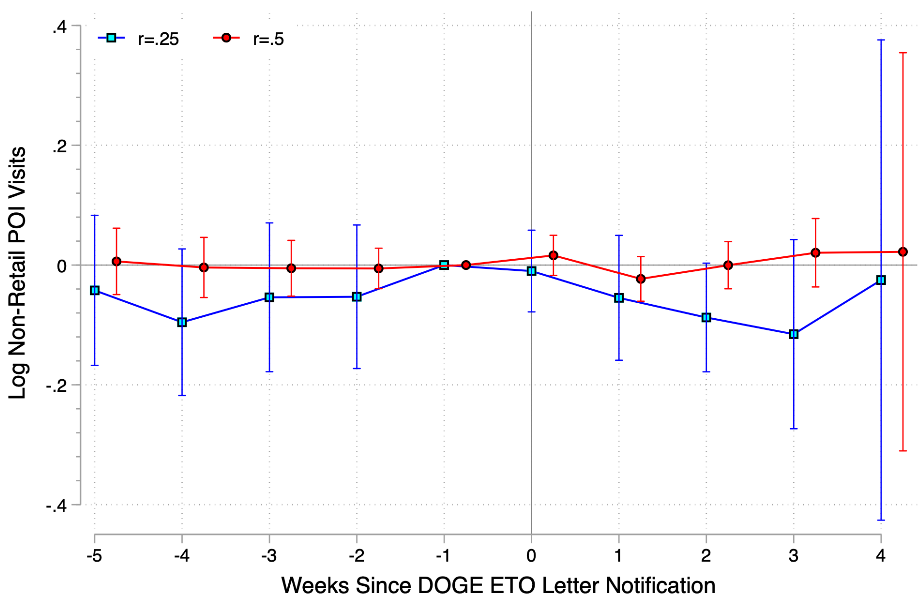
Notes: The map plots 0.5-mile rings around DOGE terminated leases that are included in the early-vs-late DiD estimation given by (C.1). Darker colors indicate rings defined by later termination cohorts, and POIs inside the darker circles serve as control group for POIs in proximity to earlier terminations, which are contained in the lightly colored circles. We restrict our sample to rings corresponding to leases terminated prior to the wave of rescissions occurring after mid-March, 2025. Orange points indicate the locations of retail POIs in the foot traffic analysis sample. Due to commercial zoning rules in Washington, D.C. most retail establishments are located on avenues, which results in clustering of retail POIs along the diagonal street grids radiating outwards from the central business district.

FIGURE C.4. Effects of Terminated Federal Leases on Foot Traffic: CSDID estimates

A. Retail POI Visits



B. Non-Retail POI Visits



Notes: We plot event study coefficients from estimating the early vs. late ring difference-in-differences specification (C.1) applied to the Advan Research (2022) foot traffic data for points of interest (POI). We estimate the event studies using the Callaway and Sant'Anna (2021) estimator, comparing foot traffic to POIs in the vicinity of an early-terminated lease to that of POIs in the vicinity of a late-terminated lease, as visualized in the map of Figure C.3. The outcome variable in all regressions is the number of visits to a POI in a given week t. In Panel A, we restrict to retail foot traffic; Panel B instead restricts to non-retail foot traffic. The outcome variable in each regression is the log number of visits to a POI. Since we restrict to a balanced panel of POIs with strictly positive foot traffic in each week, the log transform does not omit any observations. In both panels, we test for spatial spillovers at ring radii of 0.25 and 0.5 miles. We identify POIs as retail establishments based on whether they belong to one of four retail two-digit NAICS sectors *Retail Trade* (sector codes 44–45), *Arts, Entertainment, and Recreation* (sector code 71) or *Accommodation and Food Services* (sector code 72). 95% confidence intervals obtained from clustering standard errors at the Census block group level.

We adopt the estimator proposed by Callaway and Sant'Anna (2021), with log of weekly visits as outcome. We report standard errors clustered at the block group level that allow for spatial correlation of POIs located within the same block group. We plot the estimated event study coefficients in Figure C.4. We continue to find a statistically flat pattern for the response of retail (Panel A) and non-retail (Panel B) foot traffic, with the point estimates close to zero regardless of the ring radius parameterization.

D Cumulative Abnormal Returns of GSA Tenant-Exposed REITs

In this appendix, we design and implement a cumulative abnormal return (CAR) analysis to quantify how the pricing of equities of public real estate investment trusts (REITs) with different exposure to federal tenants reacts to the initial set of DOGE ETO notifications.

D.1 Event Studies of REIT Stock Prices and Returns

We track two portfolios of publicly traded REITs: a treated group with high direct U.S. federal tenant exposure to the D.C. metro office market (DEA, CDP, JBGS, OPI) and a control group holding D.C.-area assets but with little to no direct federal leasing exposure (BXP, FRT, ELME, AVB).[1] Let $P_{i,t}$ denote the closing stock price for REIT i on day t and $t_0 = 1/30/2025$ be the day DOGE first sent cancellation notifications to landlords. We re-base each series to 100 at t_0 and aggregate within portfolios such that

$$\tilde{P}_{it} \equiv 100 \times \frac{P_{i,t}}{P_{it_0}}.$$

For the treated basket we form two indices using 2024Q4 fundamentals: (i) an NOI-weighted index, $I_t^{\text{NOI}} = \sum_{i \in \mathcal{T}} w_i^{\text{NOI}} \tilde{P}_{it}$, and (ii) a square-footage-weighted index, $I_t^{\text{SQFT}} = \sum_{i \in \mathcal{T}} w_i^{\text{SQFT}} \tilde{P}_{it}$, with $\sum_i w_i^{\cdot} = 1$ in each case. The control index is an equal-weighted mean across the control group of REITs, $I_t^{\text{CTRL}} = \frac{1}{4} \sum_{i \in \mathcal{C}} \tilde{P}_{it}$. The sample runs from January 1, 2024 to October 16, 2025. Vertical reference lines mark February 1, 2025 (DOGE policy shock) and September 15, 2025 (initial broad press coverage of federal government shutdown risk).

In Figure D.1, we document that stock prices for GSA tenant-exposed REITs react quickly and negatively around January 30, 2025, consistent with equity markets incorporating the newly announced and expected future lease cancellations. Consistent with the increased salience of

[1] We classify REIT office exposure to federal tenants in D.C. based on 2024Q4 10-K (Schedule III supplemental) filings and investor materials. The treated group of REITs are heavily concentrated in buildings leased to federal agencies or to tenants whose revenue is closely tied to federal missions, especially in and around Washington D.C. The control group consists of large listed landlords with similar exposure to property, interest rate, and equity market risk but whose cash flows are primarily driven by private tenants rather than federal leases.

vacancy risk, the depth of this price decline is larger under SQFT weighting than NOI weighting. Stock prices for the two groups of REITs trended similarly in the months leading up to the creation of DOGE, including around the November 2024 presidential election. The drop in prices around October 1, 2025, when the federal government temporarily shutdown due to a lapse in appropriations legislation suggests news-beta to Washington headlines rather than a reversal in fundamentals. Together, the patterns are consistent with the evidence presented for real estate debt markets; equity REITs concentrated in federal tenancy and D.C. office ownership incur a persistent discount relative to otherwise similar D.C. owners without federal lease dependence.

Next, we test whether the market processed the abrupt change in the federal leasing policy as an increased salience of government contract risk by investigating the cumulative abnormal return (CAR) of the aforementioned treatment and control REITs groups.[2] For each index (treated NOI-weighted, treated SQFT-weighted, control equal-weighted), we estimate a standard market model over an estimation window that runs from 360 to 90 days before the event date (from February 5, 2024 to November 1, 2024), regressing daily portfolio returns on the return of the broader stock market (S&P 500).[3] The fitted intercept and slope from this regression provide a benchmark for normal returns driven purely by usual co-movement with the market. Since the procedures are standard, we defer them to the next subsection (Appendix D.2).

Figure D.2 Panel A shows cumulative abnormal returns, obtained by summing up these daily abnormal returns from 15 trading days before to 15 days after the event, separately for the two treated indices and the control portfolio. Panel B summarizes the same information in a more compact form by reporting cumulative abnormal returns over symmetric windows of different lengths around the notification date. The two panels show a sharp and persistent repricing of federal tenant exposure at the time of the initial DOGE ETO notifications. In Panel A, both treated indices experience a large drop in cumulative abnormal returns on the event date and remain roughly 3% to 5% below pre-event levels over the subsequent 15 trading days while the control portfolio rebounds and posts positive cumulative abnormal returns over the same horizon. Panel B reinforces this pattern: cumulative abnormal returns for the treated portfolios are negative and grow in magnitude as the window widens, whereas the control portfolio records modestly positive cumulative abnormal returns. Consistent with our debt market salience results, investors mark down landlords that rely on federal leases while leaving comparable REITs without such exposure largely unaffected providing equity market evidence of repricing of government contract risk.

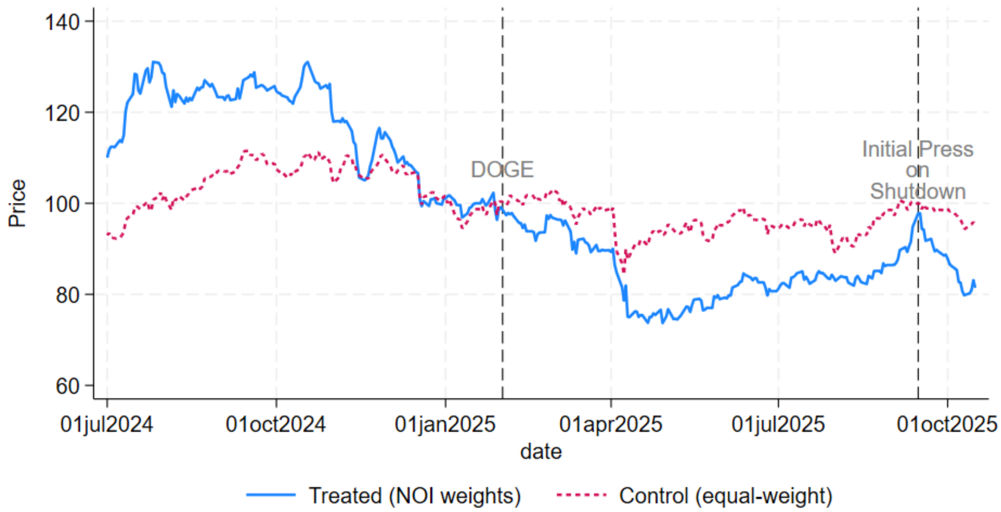
---

[2]We define the returns as a capital gain; however, our results are nearly unchanged if we instead compute total returns. All eight REITs in the sample are regular dividend payers, typically on a quarterly schedule. For example, OPI has an ex-dividend date on January 27, 2025 with payment on February 20, 2025, and DEA, CDP, and JBGS have recurring quarterly dividends. Over the relatively short event windows that we study (up to thirty trading days) the total dividend yield that falls inside a given window is on the order of tens of basis points to at most a few percent, and some windows contain no ex-dividend dates for several REITs. The large negative CARs we document for the treated indices are therefore mainly driven by price reactions.

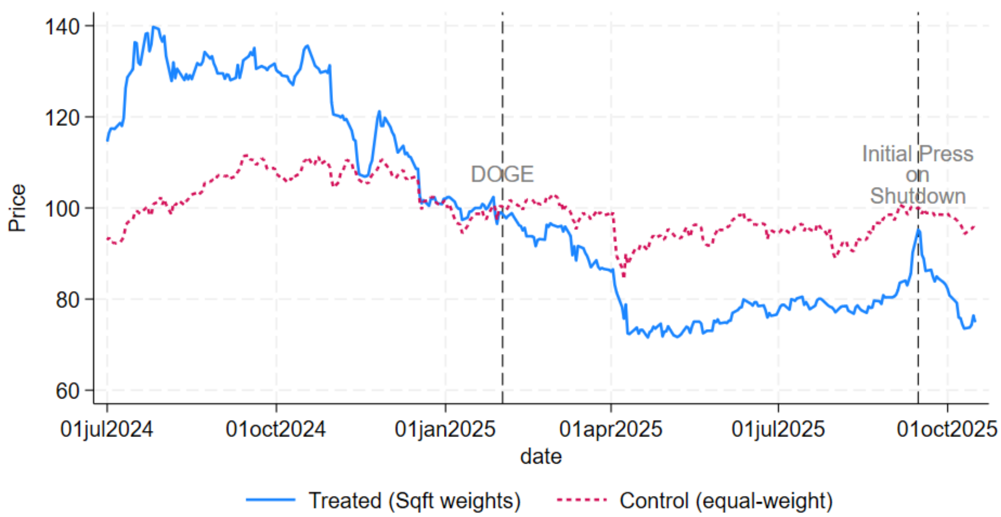
[3]We cut the normal time window to end before November 2024 to avoid any market volatility around the resolution of political uncertainty associated with the outcome of the 2024 presidential election.

FIGURE D.1. Time Series of Stock Prices for Treatment and Control Group D.C. REIT Indices

A. NOI-Weighted



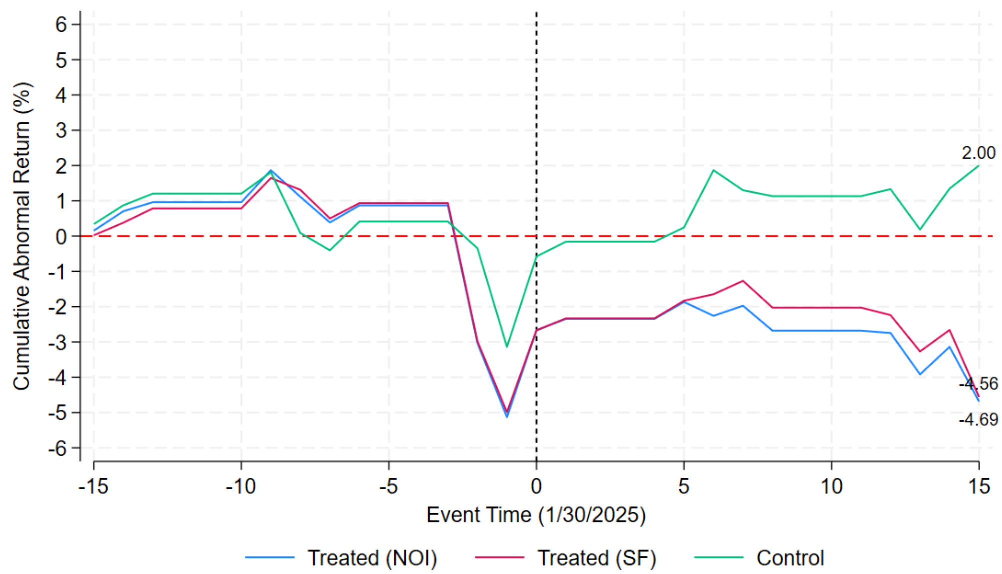
B. SQFT-Weighted



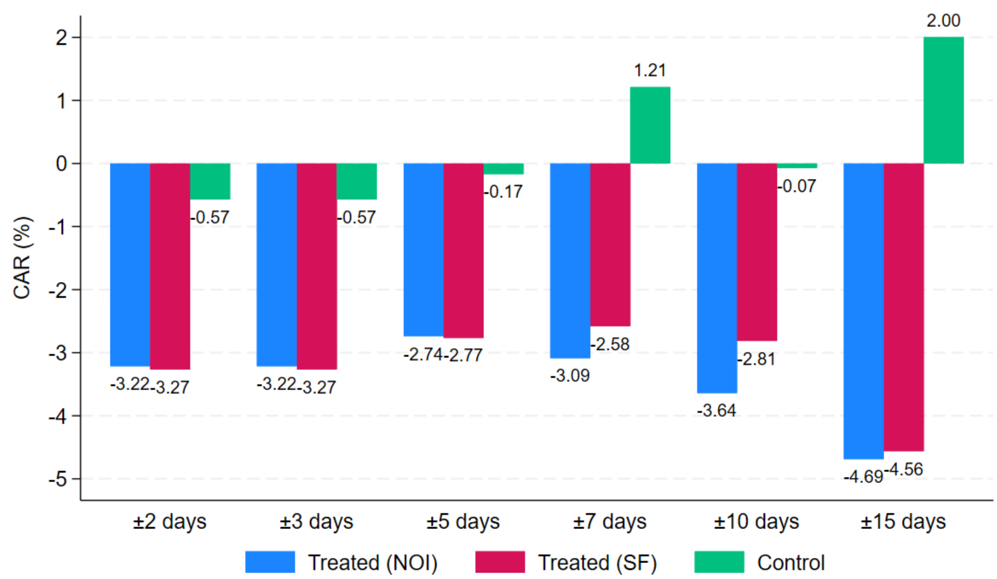
Notes: The figure plots the daily price indices for two REIT portfolios consisting of treated companies (DEA, CDP, JBGS, OPI; high D.C. office exposure) and control (BXP, FRT, ELME, AVB) REITs using prices from Jan 1, 2024 to Oct 16, 2025. Each constituent price is re-based to 100 on January 30, 2025, then aggregated. The treated portfolio yields two indices: NOI-weighted (Panel A) and SQFT-weighted (Panel B) based on the 2024Q4 10-K Schedule III information, while the control portfolio is equal-weighted. Vertical reference lines mark February 1, 2025 (DOGE-induced federal lease policy change) and September 15, 2025 (initial press coverage on the 2025 U.S. government shutdown).

FIGURE D.2. Cumulative Abnormal Returns for D.C. REITs

A. Cumulative Abnormal Returns for a Symmetric 15-day Window



B.Cumulative Abnormal Return at Various Horizons



Notes: The figure plots the cumulative abnormal return (CAR) analysis for two REIT portfolios with treated (DEA, CDP, JBGS, OPI; high D.C. office exposure) and control (BXP, FRT, ELME, AVB) groups using prices from February 5, 2024 to November 1, 2024 as the normal market return window and the symmetric windows of 2, 3, 5, 7, 10, and 15 trading days. Panel A plots the daily CARs with a symmetric 15-day window around January 30, 2025. Panel B plots the CARs for the aforementioned horizons.

D.2 Constructing Abnormal Returns: Steps

Define the daily capital gain for the stock price of REIT i as:

$$r_{i,t} = \ln\left(\frac{P_{i,t}}{P_{i,t-1}}\right),$$

where $P_{i,t}$ is the price of REIT i on day t.

Then we can construct portfolio returns for the treated set of REITs as a weighted average of the capital gains, based on either NOI or square footage (SQFT) weights.

$$r_{T,\text{NOI},t} = \sum_{i\in\mathcal{T}} w_i^{\text{NOI}} r_{i,t}, \qquad r_{T,\text{SQFT},t} = \sum_{i\in\mathcal{T}} w_i^{\text{SQFT}} r_{i,t},$$

where $\mathcal{T} = \{\text{DEA}, \text{CDP}, \text{JBGS}, \text{OPI}\}$ is the treated set of REITs with GSA tenant exposure, and $\sum_{i\in\mathcal{T}} w_i^{\text{NOI}} = \sum_{i\in\mathcal{T}} w_i^{\text{SQFT}} = 1$.

Similarly, the return on the portfolio of control REITs is an equal-weighted average of the capital gains for equities of REITs with little to no GSA tenant exposure:

$$r_{C,t} = \frac{1}{4}\left(r_{\text{BXP},t} + r_{\text{FRT},t} + r_{\text{ELME},t} + r_{\text{AVB},t}\right).$$

Let $r_{M,t}$ be the log market return on the S&P 500 on day t. For each portfolio $p \in \{T, \text{NOI}; T, \text{SQFT}; C\}$, we estimate the market model over the pre-event estimation window $t \in [T_0, T_1] = [-360, -90]$:

$$r_{p,t} = \alpha_p + \beta_p r_{M,t} + \varepsilon_{p,t}, \quad t \in [T_0, T_1].$$

Using the estimated coefficients $(\hat{\alpha}_p, \hat{\beta}_p)$, we define the "normal" return and abnormal return for all event days t:

$$\widehat{r}_{p,t} = \hat{\alpha}_p + \hat{\beta}_p r_{M,t}, \qquad AR_{p,t} = r_{p,t} - \widehat{r}_{p,t}.$$

Here, event time is defined so that $t = 0$ is the event date of initial DOGE ETO notifications on January 30, 2025.

We compute standard errors for the estimate coefficients and corresponding abnormal returns (AR) using Newey-West standard errors with four lags to adjust for serial correlation and heteroskedasticity (Newey and West, 1987). We select the minimum possible lag order such that the estimator for the covariance matrix is consistent, or `floor`$(T^{1/4})$ = `floor`$(270^{1/4})$ = 4.

For a symmetric event window of half-width h days,

$$\mathcal{W}(h) = \{-h, -h+1, \ldots, -1, 0, 1, \ldots, h-1, h\},$$

the cumulative abnormal return (CAR) for portfolio p is

$$CAR_p(h) = 100 \sum_{t \in \mathcal{W}(h)} AR_{p,t} = 100 \sum_{t=-h}^{h} AR_{p,t},$$

for $h \in \{2, 3, 5, 7, 10, 15\}$.

E Loss Simulation Procedures

This appendix provides details about how we calibrate the simulation exercises in Section 7 based on our arbitrage pricing framework for valuing lease contingencies.

E.1 Hedonic Regression

TABLE E.1. Hedonic Regression Results

Variable	**Coefficient**	**Std. Error**	***t*-Stat**	***p*-Value**
Construction Year	0.0117***	0.0026	4.53	0.000
Rentable Area (1K sqft)	-0.0020***	0.00051	-3.85	0.000
Occupancy Rate	-0.0322	0.0341	-0.94	0.347
Primary Tenant Share	-0.0029	0.0033	-0.88	0.381
Prior Debt Service Coverage	0.1044	0.1048	1.00	0.321
Prior-Year Occupancy	-0.0172	0.0266	-0.65	0.517
Prior-Year NOI ($1K)	0.00015***	0.00004	4.16	0.000
Intercept	-15.499***	5.1410	-3.01	0.003
# Properties	139			
R^2	0.505			

Notes: Dependent variable is the natural logarithm of property value. Rentable area and prior-year NOI are expressed in thousands. Regression results are estimated using the sample of Washington, D.C. Trepp properties with non-missing covariates. Robust standard errors are clustered at the property level. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

We estimate a hedonic regression model to examine the determinants of the log value of the property appraised value at securitization, using property-level characteristics with standard errors clustered at the property level with the deal pool fixed effects which accounts for the issuance of the debt claim on the property. The specification includes covariates such as property construction year (`propyear`), rentable area in square feet (in thousands, `rentareak`), occupancy rate (`occrate`), primary tenant share (`lesseepct1`), and lagged financial performance variables including the debt service coverage ratio (`priorfydscr`), prior-year occupancy (`priorfyocc`), and prior-year net operating income (in thousands, `priorfnoik`).[4] The model is estimated over 139 Washington, D.C. properties with non-missing covariates in Trepp.

Among the explanatory variables, `propyear`, `rentareak`, and `priorfnoik` are highly statistically significant. Newer properties and those with stronger prior-year NOI performance are associated with higher appraised values at securitization, while properties with larger rentable areas are systematically associated with lower appraised values, suggesting that valuation discounts are applied to larger assets in this market. By contrast, `lesseepct1`, `priorfydscr`, `occrate`, and `priorfyocc` are not statistically significant at conventional levels, suggesting that tenant concentration, contemporaneous occupancy, and lagged operating metrics outside of NOI are not systematically priced into appraised valuations in this sample.

E.2 Hedonic Model Goodness-of-Fit Tests

Coefficient of Determination & Root Mean Square Error As part of our robustness and goodness-of-fit checks, we assess the in-sample explanatory power of the hedonic regression model in (7.2). The model produces an adjusted R-squared of 47.8%; this indicates that the specification explains a meaningful share of the variation in securitized property valuations, though a certain portion remains unexplained. Additionally, the root mean squared error (RMSE) is estimated at 1.104, reflecting the average magnitude of the residuals. Taken together, the adjusted R^2 and RMSE suggest that the model provides a reasonable approximation of observed appraised values within the estimation sample, while leaving scope for additional unobserved factors that drive property-level valuations.

Mean Absolute Percentage Error To further assess model accuracy, we compute the Mean Absolute Percentage Error (MAPE), which measures the average absolute deviation between predicted and observed securitized appraised values as a percentage of the actual value. The resulting distribution of MAPE values, calculated over 139 properties, exhibits a mean of approximately 1.84 and a standard deviation of 5.42. The median MAPE is substantially lower, at 0.57, indicating that the majority of predictions are relatively close to the observed values. However, the presence of extreme outliers is evident in the right tail of the distribution, with the 90th and 99th percentiles reaching values of 2.98 and 36.48, respectively. This right-skewed distribution is further confirmed by a skewness of 5.80 and a kurtosis of 37.17, suggesting that while the model performs well for most properties, a small subset of properties experience unusually large prediction errors.

Likelihood-Ratio Test vs. Restricted FE Model To benchmark the explanatory power of the full hedonic specification, we estimate a restricted model that includes only fixed effects, absorbing variation at the deal level (`dosname`) and clustering standard errors at the property level. The restricted model (fixed effects only) explains none of the variance, while the full

[4] Variables in parentheses refer to the corresponding pneumonic in the Trepp data.

model improves this to 47.8%, leading to a 47.8 percentage points incremental R-squared from incorporating structural covariates such as construction year, rentable area, tenant concentration, and lagged financial performance. This improvement in model fit indicates that the hedonic variables are economically meaningful and provide explanatory power beyond what is captured by deal fixed effects alone.

E.3 Simulation Dynamics

We simulate the evolution of property values under early termination risk using stochastic processes that incorporate both continuous market volatility and discrete policy shocks.

TABLE E.2. Simulation Model Parameters

Symbol	Parameter	Value	Description
μ	Drift rate	0.020	Annual baseline growth in property value
σ	Valuation Volatility	0.100	Annualized standard deviation of value (market risk)
β_{DOGE}	Jump size	0.055	Average proportional loss upon ETO notification (DOGE)
β_{ETO}	Jump size	0.015	Average proportional loss for ETO non-notified (ETO)
β_{SPILL}	Jump size	0.101	Average proportional loss for non-GSA leases (SPILL)
λ	Jump intensity	0.150	Annual probability of ETO execution (Poisson arrival rate)
T	Simulation horizon	5	Time period in years over which values evolve
N	Monte Carlo iterations	50,000	Number of random draws used per simulation run
r	Discount rate	0.070	Applied rate for present value conversion
g	Growth rate	0.010	Long-run income growth used in capitalization formula

We conduct the simulation over a five-year horizon ($T = 5$) and repeatedly compute valuations and losses across 50,000 independent Monte Carlo draws. We calibrate the model parameters in Table E.2 to reflect empirical features of commercial real estate markets and observed characteristics of federal lease terminations. We select a five-year horizon to match the typical length of the soft term for ETO-eligible leases. The drift term is set to $\mu = 0.02$, representing a modest long-run expected growth in property values under normal conditions. We set the volatility parameter to $\sigma = 0.10$, capturing annualized standard deviation in value consistent with historical CRE return data (Axios Local, 2024; Cresa, 2025).

Jump magnitudes are governed by $\beta_{DOGE} = 0.055$, indicating a 5.5% average decline in value conditional on an early termination event based on our empirical results in Section 6.2. The jump magnitudes for those that are eligible for ETO but not notified and private leases are 0.015 (β_{ETO}) and 0.101 (β_{SPILL}) respectively. We estimate β_{DOGE} and β_{ETO} from regression (5.1) with log NOI as the outcome, using the first-loss group of tranches and the set of soon-to-be ETO-eligible leases as a control group for both the DOGE-notified and already ETO-eligible groups. Our estimate for β_{SPILL} originates from the spatial difference-in-differences specification (5.3) comparing private-tenant leases within a 3-mile radius of DOGE-notified properties to soon-to-be ETO-eligible leases. For each group g's jump diffusion process, we then set the jump size according to $Y_g = \log(1 - \beta_g)$.[5]

We set the jump intensity to $\lambda = 0.15$, reflecting an average annualized 15% chance of ETO execution. This parameter is calibrated to DOGE lease records and monthly termination frequencies reported in Panels A and B of Figure 10. The termination frequency is also corroborated by industry analysis at the height of DOGE-listed cancellations (CoStar News, 2025b).[6] The discount rate $r = 0.07$ and long-run growth rate $g = 0.01$ follow common assumptions in asset pricing literature and align with stabilized capitalization rates observed in recent D.C. office property transactions (DC Office of Tax and Revenue, 2025).

We model the evolution of property value V_t over a five-year horizon ($T = 5$) using stochastic differential equations tailored to each of three scenarios: ETO-notified, ETO-eligible, and private-tenant spillovers.

Scenario 1: ETO-Notified

$$dV_t = \mu V_t^- dt + \sigma V_t^- dW_t + V_t^- (Y_{DOGE} - 1)\, dC_t \tag{E.1}$$

Here, C_t is a Poisson process with intensity λ, and Y_{DOGE} represents the multiplicative impact of jumps. This specification captures the combination of continuous market volatility with sudden discontinuous declines triggered by federal lease cancellations.

Scenario 2: ETO-Eligible without Notification

$$dV_t = \mu V_t^- dt + \sigma V_t^- dW_t + V_t^- (Y_{ETO} - 1)\, dC_t \tag{E.2}$$

In this case, the structure remains the same as in Scenario 1 but with a smaller jump component, Y_{ETO}, reflecting the lower but still elevated termination risk faced by properties that are eligible but not formally notified.

Scenario 3: Private-Tenant Spillovers

$$dV_t = \mu V_t^- dt + \sigma V_t^- dW_t + V_t^- (Y_{SPILL} - 1)\, dC_t \tag{E.3}$$

Spillover properties are not directly subject to federal termination (private lease) but may experience correlated losses through demand reductions, diminished local activity, and adverse market signaling. We show in our empirical results that the jump component Y_{SPILL} is likely

[5]By calibrating the model to an annual frequency using estimates from monthly commercial loan data, we implicitly assume that properties experiencing cash flow declines do not subsequently recover. This is borne out by the persistence of the magnitude of the drop in NOI in our event study evidence.

[6]Setting $\lambda = 0.15$ implies that the average waiting time for the next lease cancellation shock is 1/0.15 = 6.67 years. This exceeds the median soft term length of five years.

driven by a combination of adverse market signaling and production externalities to the nearby tenants which provide services to the government agencies notified by DOGE. This jump-diffusion process thus captures the indirect yet economically meaningful disruptions that diffuse into nearby private leases.

In all cases, the terminal value V_T is capitalized using the Gordon growth formula:

$$\widetilde{V}_i = \frac{V_T}{r - g}, \qquad L_i = (\widehat{V}_i - \widetilde{V}_i)$$

where losses are defined as the gap between the capitalized terminal value and the benchmark fitted value from the hedonic regression $\widehat{V}_i$. We aggregate losses L_i over $N = 50{,}000$ Monte Carlo simulations to estimate risk distributions.